Identifying the Employment Effect of Invoking and Changing the Minimum Wage: A Spatial Analysis of the UK.*

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Abstract

This paper assesses the impact of the National Minimum Wage (NMW) on employment in the UK over the 1999-2010 period explicitly modeling the effect of the 2008-10 recession. Identification of invoking a NMW is possible by reference to a pre-period (prior to 1999) without a NMW. Separate identification of the effect of incremental changes in the NMW is facilitated by variation in the bite of the NMW across local labour markets with the use of the 'incremental differences-in-differences' (IDiD) estimator. We address the issues of: the possible endogeneity of the Kaitz Index; the dynamic structure of employment rate changes controlling for regional demand side shocks induced by the recession and explicitly take account of the spatial dependence of local labour markets. We conclude that there is a small negative effect of the MW introduction but no discernible effect from the uprating of the NMW on a yearly basis.

Keywords: Minimum Wage, Employment, Spatial dependence

JEL codes: J21, J38, R10

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1 Introduction

The introduction of a minimum wage (MW) could have important implications for employment levels in an economy. Likewise the uprating or changing of a MW on an annual basis could also have separate incremental effects on employment levels in the economy. Up to now, the literature has not distinguished between the imposition of a new MW and its uprating, simply because in most countries we do not observe the pre-period prior to the introduction of a MW to set a benchmark from which to measure the effect of the introduction. The introduction of a new National Minimum Wage (NMW) in Britain in 1999 and its subsequent annual uprating provide a unique opportunity to distinguish between these two effects.

The important concern of what changes should be made to a MW in times of recession, when most wages are declining in real terms, is a current and pressing problem. The problem is compounded by the consideration of what effects the MW itself may have on employment during the biggest recession since the 1930s. Since inception, the UK NMW has been administered on a national basis, with both adult and youth rates applying to all parts of the country. However, the issue of whether a MW adequately reflects regional variation in the regional cost of living, the relative balance of industrial regional growth and the growth and variation in regional productivity is questionable. Clearly longstanding geographic variation in wage rates across the UK does indeed have consequences for the bite of the NMW in different areas. As Stewart (2002) points out, the NMW reaches further up the wage distribution in certain parts of the country than in others. This study makes use of both this geographical variation and the variation in the real level that the NMW has been set at over time in order to see how changes in the local area NMW incidence over several years of the minimum wage's existence are correlated with changes in local area performance. We are also concerned that all of our geographic locations are not independent labour markets but interconnected contiguous markets. The very fact that the comparative prosperity of the South East of the country is conditioned by the economic gravity that is induced by proximity to London means that we should not treat local labour markets as independent observations in any statistical model. As Dube et al (2010) recognize, the likely consequences of erroneously doing so induces an underreporting of the standard error of the estimates and hence make mistaken positive or negative inferences on the relationship between the MW and employment.

This paper builds on that literature by examining the impact of the NMW in the UK over the period 1997-2010, comparing the period two years before its introduction with the subsequent history of the NMW and its up-ratings. This enables us to provide an additional insight by distinguishing between effects in a NMW policy off period compared to each incremental up-rating of the NMW in subsequent years. Hence instead of using a simply policy on - policy off, difference-in-difference model, we examine a model in which each year's change in the NMW is considered as a separate interaction effect. This 'Incremental Diff-in-Diff' (IDiD) estimator (introduced in Dolton et al. 2011) is a logical corollary of the econometric model suggested by Wooldridge (2002) and Bertrand et al. (2004) in the sense that it introduces a yearly interaction for each up-rating of the NMW so that we may gauge the impact of each change in the NMW. We use this IDiD procedure to evaluate the year on year impact of the uprating of the NMW on employment.

Most existing UK studies have focused on the impact of the introduction of the NMW, finding, broadly, that the aggregate employment effects of the introduction were zero or small and positive, (Stewart, 2002, 2004a, 2004b, Dolton et al 2011a). Arguably this counter-intuitive employment effect could be due to the fact that any long-run effects have not been captured by previous studies or that the problem of identifying the introduction of the NMW has not been separated from the effects of the annual uprating of the NMW. Clearly the overall effect of having a minimum wage in the labour market may induce a long run impact whereas small changes in the uprating of the level of the NMW in any given year may induce short run adjustment effects. In this paper we take a medium to long run look at the impact of the NMW in the UK and its up-ratings and try to assess whether these two separate processes may have had a differential impact across heterogeneous geographical areas.

There is a large literature on the employment effects of a Minimum Wage (see Brown et al. (1982), Card and Krueger (1995), Brown (1999), Neumark and Wascher (2009) for extensive reviews of the literature). In recent years there is a growing literature which attempts to identify the effects of a MW on employment by using geographical variation in the bite of the MW in spatially separated markets (See Card (1992), Neumark and Washer (1992, 2007), Card and Krueger (1994, 2000), Burkhauser et al (2000), Dube et al. (2007, 2010) or Baskaya and Rubenstein (2011) for the United States, Baker et al (1999) for Canada, Bosch and Manacorda (2010) for Mexico, Stewart (2002, 2004a, 2004b), and Dolton et al. (2008, 2011a) for the UK). This literature has not concerned itself with what happens to employment effects of the MW in times of macro-economic recession. This paper focuses on the modern era in the UK from 1997-2010 with the introduction of the NMW in 1999 and then leading into the current recession of 2008-2010. Hence we focus on the important question of what is the impact of the MW in an era when the economy is contracting, unemployment is rising and real incomes are falling for many people in the economy. We do this explicitly by controlling for regional demand shocks using data on Gross Value Added which is a direct measure of the level and shocks to economic activity over time at a regional level.

A second feature of nearly all the literature on the MW to date which uses geographical variation to identify the impact of the MW is that it has made the assumption that the geographical units of observation are geographically separate and unrelated to one another[1]. This assumption is unwarranted for many important reasons – we focus on just two motivations. Firstly, in reality a job vacancy is never posted with the condition that nobody outside the immediate geographical vicinity need apply. Clearly being able to travel to the job location is the problem of the individual and the resulting commute is never considered in whether someone gets the job. This means that labour markets are not neat, independent units of observation which bear no relation to one another. Economists frequently talk about local labour markets with the notion that if one considers a specific geographical area where most people both live and work in the same location then we can consider the modeling of all such areas as a set of independent, unrelated observations for our data. In reality, such a notion is false as all geographical areas have people who live in them but work in other locations. This pattern of commuting is then, in some sense, the realized form of all the subtle interrelationships between different geographical locations. A second flaw with treating such geographical units as independent is that spatially located phenomena like plant closures have an effect not just in the geographical location it occurred in but also in the immediate neighbouring areas. The degree of contiguity of neighbouring locations is therefore an important factor in the spread of unemployment, poverty, wage rises and other labour market phenomena. The extent of spillover effects from one location to another will depend on transport links, the spatial distribution of related industries and many other factors. It is well known that if we model an econometric relationship under the mistaken assumption that the units of observation are independent of one another (spherical) – when in reality they are not - then we may get biased and inconsistent estimates of the resulting economic relationships. This means that if we estimate a model of the effects of the MW using geographical data under the assumption of non-spatially related units, when they are indeed spatially related, then we will get estimates of the effects which are biased, larger than they should be and also more statistically significant than they ought to be. Hence the assumption of spatial independence is a very important one in this context which should be tested (and the appropriate estimation procedure adopted if there is indeed a spatial dependence).

An important problem that has been a preoccupation with papers in this literature is how to capture the autoregressive process of employment determination. Many papers have adopted the practice of attempting to control for this by using unemployment or various lags of employment. (See Neumark and Wascher, 1994 or Burkhauser et al 2000). Clearly such variables are endogenous to the employment dependent variable. To overcome this problem we adopt an Arellano Bond

[1] One exception is the study by Dube at al. (2010) who consider cross state border spillovers of the MW in the fast food industry in the US.

system GMM IV estimator which explicitly controls for the lagged value of employment by a carefully constructed set of lagged values as IVs. This paper is the first in the MW debate (to our knowledge) to adopt this more robust consistent estimation strategy of a dynamic panel model.

A difficult issue that this paper attempts to tackle head-on is the problem of the appropriate way to address the issue of the potential endogeneity of the controlling MW regressor – typically the Kaitz index. Early papers in the literature just assumed that the wage regressor in the model was exogenous. More recent authors recognize the problem and use a variety of IV type strategies to solve it. (See Dube, 2010, Dolton et al 2011, Baskaya and Rubenstien 2011) Our work is less open to this criticism as we operate with data from a country in which the MW is set as a NMW – and therefore, from the perspective of any specific geographical unit the ration of the NMW to the median wage in that geography is unlikely to be related to the unobserved heterogeneity in employment at that geography. Nonetheless we wish to allow for the possible endogeneity of this regressor due to the potential influence of the denominator in the Kaitz index relating to median wage movements. Therefore we take seriously the issue of the potential endogeneity of the Kaitz index and the approach we adopt is novel in that it seeks to use an spatial IV strategy to identify this effect. Specifically this means using the nature of the spatial relationships of the geographical units to deduce an IV for the Kaitz index which is based on the neighbouring and related areas respectively. We report that this meets the first stage criterion of a good IV in the sense that it is highly correlated with the Kaitz in the specific region.[2] Of course, for the second stage, we cannot prove that the instrumented value of the Kaitz for any location is not correlated with the unobservables but we adopt, the not unreasonable assumption, that the possible unobserved heterogeneity in a region may not be correlated with the observed values of a weighted average of the actual Kaitz index in the neighbouring areas.

Crucially in this paper we examine whether the definition of the geographical unit used for the analysis matters. Since the definition of what constitutes a 'local labour market' in Great Britain is still open to discussion, the analysis is undertaken at three different levels of geographical aggregation. As in Stewart (2002), the data can be divided into 140 areas comprising Unitary Authorities and Counties. However, the same analysis can be done using 406 Unitary Authorities and Districts. We also look at how the results change if we use the definition of 67% of people living and working in the same geography to capture a local labour market, as now used by the UK national statistics office to define a "travel to work area" (TTWA). We remain agnostic as to what

[2] Formally the use of a spatial IV may only lead to consistent estimates under certain circumstances. (see Fingleton and Gallo 2008 and Hoxby 2004).

the correct definition of a 'local labour market' is and let the data tell us whether such definitional difficulties matter.

A secondary goal of this paper is that it attempts to set the different estimates in the literature in context as our econometric estimation model is more general than previously used. In this respect we introduce each of our contributions in a stepwise fashion to establish that we can generate the results of earlier papers – but that when we add important considerations like the spatial dependence, regional aggregate demand shocks, lagged dependent variable regressors and other endogenous regressors then we find clearer – less contradictory results. Therefore, to some extent, our findings can nest those of earlier contributions. Hence we examine the robustness of our results with regard to the specification issues associated with: spatial dynamic specification to incorporate the lagged effects of previous employment on current employment, and time and interaction effects. In this testing of robustness we suggest that much of the previous literature is sometimes presented as if the results are in stark contrast to each other. Our take on this literature is that it often estimates fundamentally different parameters and that this explains a large degree of the differences in results.

To summarize; this work contributes to the literature in five ways. Firstly we separately identify the employment effect of having a MW from the possible effects of uprating the MW. Secondly, unlike the literature, we treat the geographical units of statistical analysis as being spatially related (rather than independent). Thirdly, we explicitly take account of the current recession by direct consideration of the role of shocks to aggregate demand. Fourthly, we directly tackle the issue of the dynamic nature of employment process (by considering the autoregressive structure of employment using Arellano-Bond type IV estimation in a system GMM context), and finally we examine the possibility of the endogeneity of the Kaitz index as a proxy for the wage effect. These advances provide a fundamentally new insight into the effect of the MW on employment.

The paper is organised as follows. Section 2 describes the datasets used and the characteristics of the data and contains a description of the main areas of novel interest in terms of the spatial nature of the data. Section 3 outlines the econometric methodology and identification for the analysis and presents the main results of the analysis. Section 4 concludes.

2 Data

The central idea of this paper is to see whether geographic variation in the “bite” of the minimum wage is associated with geographic variation in employment. Geographical variation in wages in the UK is exploited in order to evaluate the impact of the NMW on employment at the local level. The data used in this study are drawn primarily from three sources. Data on earnings, and a restricted number of covariates all disaggregated by geography is provided by the New Earnings Survey (NES) from 1997 to 2003 and by the Annual Survey of Hours and Earnings (ASHE), which replaced the NES in 2004. In both surveys, conducted in April of each year, employers are asked to provide information on hours and earnings of the selected employees. The geographic information collected for the full sample period used in the paper is based on workplace rather than residence. In what follows we describe the data. We provide details of our data and its limitations and the technical properties of our key variables in our Data Appendix A. In this appendix we also discuss the alternative geographical units of analysis.

The geographic variation in wages will reflect the demographic and industrial composition of each local labour market. The changing industrial composition of an area and the extent to which industries are low and high paying will affect the changing incidence of the minimum wage working in a locality. Likewise the skill, age, gender and sector composition of the local workforce will be important factors. To a certain extent we can control for variation in these influences with a set of time varying local labour market control variables, drawn from either ASHE or matched in from complementary Labour Force Survey (LFS) data. However, the choice of what constitutes a local labour market is open to discussion, therefore the analysis is conducted at three different levels of aggregation. We perform our estimation separately, on the level of 140 Unitary Authorities and Counties (WAREA), on the level of 406 Unitary Authorities and Districts and finally on the level of 138 travel to work areas (TTWA).

Since the choice of the unit of geographical analysis is central to this paper as it serves to highlight the main differences between the different levels of analysis’. Specifically the: 140 level is borne of administrative areas and consists of large counties and separate conurbations; the 406 areas is based on a much finer level of detail of administrative sub-areas consisting of constitiuent parts of the main 140 areas which are much smaller and compact; and the 138 TTWAs are defined explicitly using the definition of a threshold of the fraction who live and work in the same area. This underlying difference in our areas will be reflected in the qualitative difference in our results. Specifically one would expect that where our geographical units are defined using commuting patterns then one might derive less distinct estimated coefficients from spatial models compared to when one uses geographical units which are defined by administrative regions. We perform our analysis for each of the three levels of geography as an important robustness test as the nature of the cross section units are fundamentally different in the case of our TTWA, WAREA or WLA areas. Since the economic geography literature, is to our knowledge, largely silent on what is the correct choice of the level of geographical analysis then we repeat our analysis for all three levels. Obviously, in our analysis the standard errors will be smaller for the 406 geography than the 138 and 140 geographies but the 406 sample has the potential drawback of having computed employment rates (for example) which are based on smaller sample sizes. We are largely reassured by the fact that our results are very comparable across the three geographies.

The employment variable

We then match local area employment data from the LFS with the minimum wage covariates generated from ASHE. There is an important feature of the timing of data collection which we exploit in order to try and make sure that our employment variable is measured after the up-rating of the NMW. The ASHE and NES estimates for hourly earnings and therefore the minimum wage variables used in this paper are recorded in April of each year. Since the minimum wage was first introduced in April 1999 but then up-rated in October of each following year, the NMW variables are therefore generally recorded six months after each NMW up-rating.[3]

Data on employment at these levels of aggregation derived from the LFS are available via NOMIS for yearly data for 1997 and 1998. For the period 1999 to 2005 we use employment rates calculated from the quarterly LFS local area data. For the years 2006 to 2010 we use the quarterly LFS Special License data to calculate the employment rate.

Measuring the National Minimum Wage

The most widely used variable to measure the level of the NMW in the literature is the Kaitz index, defined as the ratio of the minimum wage relative to some measure of the average wage. We use the median wage in our study. The closer the Kaitz index to unity the “tougher” the bite of minimum wage legislation in any area. However the denominator can be influenced by factors other than the level of the NMW and so the median wage is arguably more endogenous in an employment regression. For example, a positive correlation between the employment rate and the median wage might be generated by an exogenous labour demand shift. This will create a negative correlation between the Kaitz index and the employment rate. Although we have used alternative measures of

[3] There are however two exceptions that are described more in detail in the Appendix A, section “Definition of key variables”.

the MW in previous work (see Dolton et al. 2011a where the fraction of people at or below the NMW and the Spike are used) we did not find that this alternative definitions of the MW variable made any qualitative difference to our conclusions. Our estimations with the other possible variables to represent the MW (not reported here) have not shown different conclusions.

Modelling Conditional and Compositional Covariates

The geographical heterogeneity of areas and localities in the UK is well known. Our analysis attempts to condition out for this spatial variability by using a vector of observed (derived) covariates. Explicitly we control for: the demographic age structure of each population (using average age, age squared and age cubed); the level of human capital in each area using the fraction of those qualified to degree standard (NVQ level 4 or above in each geography); the fraction of each population of working age which is female; and the compositional industrial structure (Duranton and Overman 2005) – specifically the fraction of who work in the public sector. The final variable requires a brief elaboration.

There is some considerable debate in the UK as to the extent to which the size of public sector 'crowds out' the private sector (Faggio and Overman 2012). There is also a considerable debate on regional inequality and the so called 'North-South' divide (Smith 1994). The thinking of the free market economists is that a vibrant growing economy needs an expanding private sector and that a large public sector gets in the way of such potential growth. This view predominated in the Thatcher era (1979-1987) and now has common currency in the present coalition government (2010-15) – but this was not the dominant view in the era leading up the NMW (1987-2010). Conversely, it is clear that the multiplier effects of having a large local public sector are obvious. It was with this aim that many of the core government departments were moved out of London to the regions in the 1980s and 1990s[4]. These are the forces which have shaped the development of the regional economies of the UK over the last 20 years. We try to take account of these changes by controlling for the fraction of each local labour force working in the public sector.

Modelling the spatial nature of labour markets

One of the main innovations in this paper is our attempt to capture the interconnectedness of local labour markets. The importance of the spatial nature of labour markets is now becoming more widely recognized and exploited in the work of Patacchini and Zenou (2007), Moretti (2011) and others. More specifically the suggestion that commuting patterns in UK are a way of representing

[4] Specifically the Department of Health moved to Leeds, the Department of Work and Pensions to Sheffield, the Department of Social Security to Newcastle, the Driver and Vehicle Licensing Authority, DVLA to Swansea

this spatial interconnectedness are being used by others recently (e.g Manning and Petrongolo 2011). Where a given 'local labour market' begins and ends and the extent of interconnectedness of spatially located areas will depend on a multitude of factors, including: distances, physical geographical features like rivers (Hoxby 2000), lakes and mountains, rail networks, (Gibbons and Machin 2006) bus links, the availability of major arterial roads, house prices (Gibbons and Machin 2003), commuting patterns (Rouwendal (1998), school quality, council tax levels, crime rates (Gibbons and Machin 2008) and the provision of amenities, to name but a few. In some sense it is impossible to observe all these factors in determining how interconnected each labour market is to every other labour market. To a degree all these influences on the spatial nature of the location decision of where one lives and works are determined by a large number of unobservables[5]. Our approach to this problem is to take the observed pattern of commuting behaviour as the empirical 'reduced form' of all these influences which we cannot possibly observe. In some sense, if the degree of interconnectedness of labour markets is, *de facto* the actual propensity for an individual to live in region A and work in region B aggregated up over all regions and all workers then this is what we should use in our calculations.

One concern that may be important is that the using the commuting matrix as a weights matrix may be potentially endogenous in the sense that its degree of interconnectedness for any specific location may be related directly to its level of economic activity[6]. To allay this concern we also use an alternative weights matrix based on geographic contiguity. This is a logical alternative used in the literature to model the spatial dependence (Möller and Aldahev, 2007). This measure is specifically 1 if a specific location borders another location, and 0 otherwise. We use this alternative weighting system as a robustness check on the grounds that such geographical divisions are largely historical and exogenous to current economic forces.

Measuring the recession

The second innovation of this paper is to attempt to net out of our estimations any underlying movements in aggregate demand and more importantly the large potential effects of the current recession. This has rarely been attempted – indeed to our knowledge the only research on this topic has been our own previous attempts to tackle this issue (see Dolton et al. 2011a, 2011b, 2012 and Dickens and Dolton 2010). This analysis was rudimentary in that it relied on simple dummy indicators as the presence of a recession or not. The problems associated with this when the formal

[5] There is a growing body of literature about in commuting behavior. This includes mapping it (Titheridge and Hall, 2006 Nelson and Hovgesen 2008), accounting for it in labour market search models (Rouwendal 1998) and econometrically modeling its determination (Le Sage and Pace, 2008).

[6] The whole issue of how a weighting matrix is conceived is discussed in Harris and Moffat (2011).

definition of a recession is two quarters of negative GDP growth are rehearsed in Dolton et al. (2011b). Here we adopt a more ambitious approach as we attempt to control for negative regional GDP growth shocks with a direct proxy for regional growth. Therefore we seek an (exogenous) variable which captures the depth of a recession on a regional basis but which is not endogenous to the determination of employment directly. The requirement that this variable is available on regional basis proved exacting. Arguably, the exogeneity requirement rules out, employment measures for other groups, unemployment or measures related to the claimant count[7]. We explored measures such as house prices as such data is collected at a local level. The problem with this data is that such series only mirror recessions with a significant and differential lag (of up to 5 years) which is dependent on geography. Another suggestion was to use VAT registrations of the birth and death of companies – but this data series ends on a regional basis in 2007. The variable which we use is the lagged level of Gross Valued Added on a regional basis which is available in Regional Trends. The definition of this variable is that it aggregates all firm revenues, profits and all wages on a regional basis to compute literally the gross value of goods and services in the regional economy. Hence, to all intents and purposes, this variable is a measure of GDP growth (per head) on a regional basis. This, in our view, is the closest one can get to a variable which measures in a continuous way the level of regional GDP growth changes over time and hence it is a variable which captures when negative aggregate demand shocks hit; when a recession occurs and how bad it is in different regions in different years. The obvious criticism of this variable is that it is potentially endogenous to employment levels in the sense that the wages of employed people are included in its calculation. But since the variable includes much more than this in terms of the values of goods and services produced we suggest this, rate of change of GVA variable can act as a proxy for the onset, timing, severity and duration of regional GDP growth and hence of recessions. A second response to this is that it is an advantage to have regional data on this demand shock variable as the demand level at any specific local sub-geography will not be an important constituent of this variable at the regional level.

The logic of identification from the data

The logic of our identification strategy is evident in the descriptive statistics we present in Figure 1 that highlights the temporal variation in the NMW, comparing the nominal hourly wage level of the adult NMW over time with the notional level which would have been achieved if the level of the NMW in 1999, when it was introduced was indexed to average earnings. The Figure shows how the

[7] Dube et al (2010) use private sector employment, Neumark and Wascher (1992) and many other authors use unemployment for adults. These measures are arguably endogenous.

NMW started off by being lower than the average rise in earnings and then rose more steeply than this series. Most marked is the rise in this level in both real and nominal terms since 2003. The largest rises in the NMW are in 2001, 2004 and 2006. This is mirrored in the rising level of the Kaitz Index over the same years shown in Figure 2. The principle here is that we wish to use the height of the steps due to the up rating of the MW considered over all the different locations in our sample.

Figure 3 is more instructive of our data as it shows the level of the NMW (right hand scale) plotted on the same graph as the employment rate (left hand scale) in our sample. It plots this for the 138 TTWAs at the mean of the sample with the 95% standard errors in dotted lines. Here we can see the period of 2002-06 where the boom years prior to the recession which began in 2008/9. The nature of this trend in employment needs to be picked up in the data and this is why we seek to model the underlying 'steady-state of employment' by seeking to identify the autoregressive nature of this process. Figure 4 adds to our understanding of what was happening to employment in relation to the movements in the Kaitz index (again at the average for the sample with 95% confidence intervals plotted). Here we see that – to a large extent – the upward movements of the Kaitz are mirrored by a downward shift to the employment level. Hence we would expect to find an overall negative relationship between these two key variables. Reassuringly – this is what we find – although what we set out to do is condition out the problems which besets this kind of data – namely endogeneity, demand shocks, and the nature of the underlying employment process.

[Insert Figure 1 Here]

[Insert Figure 2 Here]

[Insert Figure 3 Here]

[Insert Figure 4 Here]

3 Methodology and identification

Invoking and change of the NMW

To understand any of the estimation results relating to the impact of the NMW one must be clear about the econometric specification and which parameters we seeks to identify in the model. We begin with the most basic model and develop it.

Neumark and Wascher (1992) are among the first to utilize panel data to address the question of the impact of the MW.[8] They estimated the model:

$$E_{jt} = \alpha + \gamma T_t + J_j + \beta MW_{jt} + \pi D_{jt} + \delta X_{jt} + \varepsilon_{jt} \tag{1}$$

Where E_t is employment at time t in State j , MW_{jt} is the level of the MW (adjusted for coverage) at time t in State j, D_{jt} is a measure of aggregate labour demand (or the recession) in region j in year t, X_t is a set of controlling regressors at time t in State j, T_t is a set of year effects, and J_j is a set of State fixed effects. Fixed effect estimation identifies potential causal inferences based on changes in the regressor and regressand given the assumption that the unobserved heterogeneity across areas remains constant over time. Later Neumark and Wascher (2004) use nearly the same specification to estimate the impact of the NMW laws across countries with the slight modification that now the MW_{jt} term is similar to the Kaitz index using the ratio of the NMW in country j at time t divided by the average wage in that year[9]. Neumark and Wascher in their various papers, whether at the US State level or at the level of countries, also find a negative employment effect of the NMW.

The logical critique of this panel model is that it still suffers from potentially all the same sources of potential heterogeneity bias as the simple time series model. Indeed it could even be argued that using geographical States as the unit of observation could potentially have even more problems - if for example - one state legislature's decision to implement or change a MW is heavily influenced by the current level of unemployment or a neighbouring state's policy decision. This concern is less of a problem in the UK context as there is a national NMW rather than a state MW - in which case the actual level (and change) in the NMW is not under the control of the authorities in any particular location.

A related methodological departure focused on identification is suggested by Card (1992) and Stewart (2002) who propose a ‘structural’ econometric model consisting of two equations. The first is a form of labour demand equation which suggests that any change in the employment rate in area j is a movement along the labour demand curve which results from a change in the wage level in area j conditioning out for any shocks in aggregate labour demand, D.

$$\Delta E_j = \gamma_0 + \eta \Delta W_{jt} + \pi D_j + u_{1j} \tag{2}$$

[8] More precisely, they used US state data from 1973-1989.

[9] Usually the Kaitz index is also weighted by some measure of 'coverage' of the NMW in the sense of the fraction of the labour force that the NMW applies to.

The second equation is a form of identity suggesting that the wage increase in area j is a function of the proportion in the area who are 'low paid', P_j.

$$\Delta W_j = \alpha_1 + \lambda P_j + u_{2j} \tag{3}$$

Substituting equation (3) into equation (2) we get:

$$\Delta E_{jt} = \gamma_0 + \beta P_j + \pi D_j + u_{1j} \tag{4}$$

Where $\beta = \eta\lambda$, with λ assumed to be positive, implying that β has the same sign as η which basic economic theory would suggest is negative if the demand for labour falls as wages rise. According to Stewart (2002) the precondition for identification is that the proportion in the area that are 'low paid', P_j is a predetermined instrument for the endogenous wage change[10].

The central idea of our paper is also to see whether geographic variation in the "bite" of the minimum wage is associated with geographic variation in employment. However we also allow the effect of any treatment to vary over time, given the differential pattern of upratings that we observe in the data. This can be done by pooling over the fourteen year period and letting the treatment be the measures of the "bite" of the NMW in each area at time t, $P_{jt,}$ so that the model estimated is:

$$E_{jt} = \gamma_0 + J_j + \sum_{k=1999}^{2010} \gamma_k Y_k + \theta_0 P_{jt} + \sum_{k=1999}^{t} \theta_k^{IDID} Y_k\, P_{jk} + \pi D_{jt} + \delta X_{jt} + \varepsilon_{jt} \tag{5}$$

Where E_{jt} is a measure of area labour market performance in area j at time t, J_j are area effects, and Y_t is a set of year effects with $Y_t = 0$ for t=1997, 1998. The range k is indexed from 1999 (the year in which the NMW was introduced and subsequently up-rated). Area fixed effects are included to control for omitted variables that vary across local areas but not over time such as unmeasured economic conditions of local areas economies that give rise to persistently tight labour markets and high wages in particular areas independently of national labour market conditions. Time fixed effects control for omitted variables that are constant across local areas but evolve over time.

The Incremental Difference-in-Difference coefficients θ_t^{IDiD} on the interaction of the year dummies and the measure of the bite, capture the average effect of the up-rating of the NMW in each year, starting from the introduction of the policy in 1999 all relative to the 'off period' of 1997 and 1998, provided of course that the proportion in the area who are 'low paid', P_{jt} is a valid instrument for the endogenous wage change. The advantage of using the IDiD estimation

[10] In reality we use the Kaitz Index to act as a proxy for the wage effect of the NMW. In our previous paper Dolton et al (2011a) we explore 2 other measures of the MW and the substantive conclusions do not differ much. In our later section we also examine the possibility that the Katiz index is itself endogenous. Here we use a novel spatial IV strategy.

procedure is that it facilitates the estimation of year on year incremental effects of each year's up-rating. So even if the average effect over all years is insignificantly different from zero, this does not mean that the effect of any individual year's change in the NMW is also zero. Note that one cannot deduce the longer run effect of all the changes in the NMW by simply summing all the year-on-year IDiD coefficients.[11] The long run effect can only be measured in aggregate by using one DiD coefficient for the whole period. We therefore present both short run IDID and medium run DiD estimates in what follows.

Though we have 3 more years with observations compared to previous work in Dolton et al. (2012) our time series remains quite short. To the best of our knowledge there is no statistical method that would find strong evidence for autocorrelation of a higher order in our panel data set. An additional concern, we also already mentioned in Dolton et al. (2011), is that spatially contiguous areas lead to heteroskedastic errors. In what follows we explicitly model these spatial relations.

Basic identification issues

One important question to ask is how long will it take the introduction (or changes) in the NMW to have its full effects on employment and other economic indicators (especially since some of the variables in the data are already measured with a lag). From an empirical point of view, this raises the specification issue about including a lagged effect of the minimum wage variable in the regression. The debate on this question is still ongoing. On the one hand, employers might react relatively quickly to increases in minimum wages. Employers might even adapt before the implementation of the minimum wage. Brown et al. (1982), regarding employment, argue that: "One important consideration is the fact that plausible adjustment in employment of minimum wage workers can be accomplished simply by reducing the rate at which replacements for normal turnover are hired.", (p.496). Another reason given by the authors is that minimum wage increases are announced months before they are implemented – typically six months in the UK - therefore firms may have begun to adapt before the increase of the minimum wage come effectively into force. On the other hand, it might take time for employers to adjust factors inputs to changes in factors prices. Hamermesh (1995) points out that in the short run capital inputs might be costly to adjust. If firms adjust capital slowly following an increase of the minimum wage, the adjustments of labour input might be slowed as well. The use of a lagged minimum wage measure as well as the inclusion of fixed (area) effects in the regression also helps to decrease the possible endogeneity of

[11] This is because some additional (untestable) assumption relating to independence of effects over time would be necessary. In addition, since we use a dummy variable interaction term, rather than a normalised metric on how large each increment was then this also makes aggregation of the individual interaction term estimates difficult.

the minimum wage variable which occurs from correlation of either the proportion paid at the minimum or, in case of the Kaitz index, the minimum wage and the median wage with labour market conditions or productivity.

An issue of identification arises from the 'common trends assumption' which, in our context, is the assumption that the effect of market conditions will be the same across all geographic units in the absence of the introduction of the NMW. One way of examining this is to consider whether the employment rate has the same underlying trend across all our geographical units before the introduction of the NMW. In our case we cannot do this because the small geography LFS data which we use to construct the employment rate does not go back before 1997. However, it is possible to have a longer off-period starting from 1994 and using 95 areas, which correspond to the coding used on the NES (the National Earnings Survey which preceded the ASHE) up to 1996.[12] The results of the test give us some confidence about the internal validity of the model, being unable to reject the null of a common trend at 10% level. [13] Whilst this is no proof of the presence of common trends in our data, this gives us some confidence about the internal validity of our model for the full set of more detailed geographies.

Modelling spatial dependencies

In recent years, the econometrics literature has exhibited a growing interest in specifying spatial dependencies, or more generally, cross sectional dependencies because estimation results could be spurious if there is spatial dependency that is not considered in the model (Le Sage and Pace 2009) .

The idea is that an economic aggregate like employment in a certain region does not only depend on economic key figures in the same region but in other regions. To model these dependencies a class of spatial econometric models were developed that consider spatially autoregressive processes in the dependent variables or in the error term. The first model is often called the Spatial Lag or Spatial Autoregressive Panel Model (SARP) and the second is called the Spatial Error Panel Model (SEMP, Elhorst 2010a).

In the following we extend our model in equation (5) with spatial lags. That means in the case of spatial lags of the dependent variable:

$$E_{jt} = \rho \sum_{i=1}^{n} w_{ij} E_{it} + \gamma_0 + J_j + \sum_{i=1999}^{2010} \gamma_i Y_i + \theta_0 P_{jt} + \sum_{k=1999}^{t} \theta_k^{IDID} Y_k P_{jk} + \pi D_{jt} + \delta X_{jt} + \varepsilon_{jt} \tag{6}$$

[12] The areas comprise all existent counties, the counties abolished with the 1996 local government reform and the London boroughs. The “City of London” was deleted from the dataset due to small sample size and the Scottish Islands were excluded from the analysis because they are not present in the data across all years.

[13] For all workers (16 years to retirement) we cannot reject the null of a common trend at the 10% level (F(91, 276)=1.45) if we omit three areas, all with small sample sizes, (Scottish Borders, Gwynedd and Shropshire). However, omitting these areas from our IDiD regressions does not change our main results.

where ρ is the coefficient for the spatial lag term $\sum_{i=1}^{n} w_{ij} E_{it}$, a linear combination of values of the employment rate from regions i that are assumed to influence the observations in region j (LeSage and Pace 2009). The weights w_{ij} contain zeros if there are no spatial lags and the main diagonal of the weight matrix contains zeros since it is assumed that a region cannot have an influence on itself. Furthermore, the weight matrix used for the estimation is standardized with rows summing to unity, irrespective of information used to model regional dependencies. The weights used reflect the assumption about the relative strength of the spatial lag. In every case it is intended to identify the spatial dimension of economic or regional activity and to implement that in the model.

A simple assumption is that neighbouring or nearby regions have a greater influence on other than those that do not share a border or vertex (LeSage and Pace 2010). The contiguity weights matrix before row-standardization contains ones in case of contiguity and zeros otherwise. Like weight matrices based on distances between regions, this matrix is symmetric. That implies that e.g. region A influences region B to the same extent as region B influences region A. It is important to acknowledge that this could be a restrictive assumption for the UK where there are clearly asymmetric economic relations between economically strong regions like London and surrounding economically weaker provinces.It is this logic which induced us to use commuting patterns. These are good indicators of the intensity of regional labour market interdependencies since they summarize spatially related economic decisions and behaviour. Furthermore, commuting streams have direction – the number of people that go from their (home) region A to work in region B differs to that number of other people that go from region B to region A. Therefore, we decided to use commuting flows and to compare our results with specifications that are based on contiguity as a robustness check (see Appendix B for some more details). Hence, we use the flow of commuters from their home region i to region j where they work.[14] To form a spatial lag or a linear combination of values from the "nearby" regions, for each region j, weights w_{1j} to w_{nj} are normalized to the (row) sum of unity.

The dependent variable, E in equation (6) is both on the right and the left side of the equation. To estimate the model equation (6) has to be rearranged for all regions $j=1,\ldots,n$. Equation (6) in matrix notation is

$$\mathbf{E} = \rho \mathbf{W_T E} + \gamma_0 \boldsymbol{\iota}_{\mathbf{nT}} + \mathbf{Y_N} \boldsymbol{\gamma}_{\mathbf{t}} + \theta_0 \mathbf{P} + (\boldsymbol{\theta}_{\mathbf{t}}^{\mathbf{IDID}\prime} \mathbf{Y_N})' \mathbf{P} + \mathbf{X}\boldsymbol{\delta} + \boldsymbol{\varepsilon} \quad (7)$$

[14] We are sensitive to the possibility that our W weighting matrix may be endogenous. In the empirical estimation we also run all our analysis with an alternative 'contiguity' weighting matrix which is simply constructed as a matrix of 1s and 0s for each geographical location depending on whether the location abuts a neighbouring location (1) or not (0).

Hence $\mathbf{E}$ ($\mathbf{P}$/ $\mathbf{Y_N}$/ $\boldsymbol{\varepsilon}$) is the $nT \times 1$ vector containing the employment rates E_{jt} (the measures for the Kaitz index P_{jt} / the year effects variables Y_t with Y_t= 0 for t=1997, 1998 / the error term ε_{jt}), $\mathbf{W}_T$ is a $nT \times nT$ matrix containing all weights w_{ij} that are equal for all years. $\boldsymbol{\iota}_{\mathbf{nT}}$ is a unity vector with dimension $nT \times 1$. The $nT \times 1$ vector $\boldsymbol{\gamma}_{\mathbf{t}}$ contains the parameters for the year effects. The $nT \times k$ matrix $\mathbf{X}$ contains the k control variables including the aggregate demand shocks measure D, and the $k \times 1$ vector $\boldsymbol{\delta}$ the coefficients of the control variables δ.

Now we can solve equation (7) for $\mathbf{E}$ and get the regression equation for model considering spatial lags of the dependent variables:

$$\mathbf{E} = (1 - \rho\mathbf{W_T})^{-1}(\gamma_0\boldsymbol{\iota}_{\mathbf{nT}} + \gamma_t{}'\mathbf{Y_N} + \theta_0\mathbf{P} + (\theta_t^{\mathrm{IDID}}{}'\mathbf{Y_N})'\mathbf{P} + \mathbf{X}\boldsymbol{\delta} + \boldsymbol{\varepsilon}) \tag{8}$$

The equation for the Spatial Error Panel Model in matrix notation is

$$\mathbf{E} = \gamma_0\boldsymbol{\iota}_{\boldsymbol{nT}} + \mathbf{Y_N}\boldsymbol{\gamma}_{\mathbf{t}} + \theta_0\mathbf{P} + (\boldsymbol{\theta}_{\mathbf{t}}^{\mathbf{IDID}}{}'\mathbf{Y_N})'\mathbf{P} + \mathbf{X}\boldsymbol{\delta} + \mathbf{u} \tag{9}$$

with a spatially autoregressive process in the error term **u,**

$$\mathbf{u} = \lambda\mathbf{W}_{\boldsymbol{T}}\mathbf{u} + \boldsymbol{\varepsilon} \tag{10}$$

Finally solving (10) for **u** and implementing that in (9) leads to the SEMP model

$$\mathbf{E} = \gamma_0\boldsymbol{\iota}_{\boldsymbol{nT}} + \mathbf{Y_N}\boldsymbol{\gamma}_{\mathbf{t}} + \theta_0\mathbf{P} + (\boldsymbol{\theta}_{\mathbf{t}}^{\mathbf{IDID}}{}'\mathbf{Y_N})'\mathbf{P} + \mathbf{X}\boldsymbol{\delta} + (1 - \lambda\mathbf{W}_{\boldsymbol{T}})^{-1}\boldsymbol{\varepsilon} \tag{11}$$

Since the regression equations in (8) and (9) are non-linear in their parameters, maximum-likelihood estimation is used to estimate the parameters. We use the estimation procedure suggested by Elhorst (2010b) that includes a bias correction for both time and spatial fixed effects (for details see Lee and Yu 2010).

The question is - which model specification should be preferred - models without spatial dependencies or the models with either an autoregressive error term or the model with a spatial lag? This is crucial because misspecification would lead to biased estimates of the coefficients of interest. We therefore conduct Lagrange Multiplier Tests which show us that spatial dependencies are present and should not be neglected and there are indications that the SEMP approach should be preferred in the vast majority of specifications (all details can be found in the Appendix C) whereas a somewhat "different" spatial specification should be preferred in other specifications. However, a shortcoming of the test is that they are only able to test the SEMP and SARP, their combination as well as a somewhat spatial specification against no spatial lag.

Gibbons and Overman (2012) showed the formal equivalence of the SEMP and SARP models with the spatial Durbin model that contains spatial lags of the dependent variable and the exogenous regressors. Furthermore, they show that the reduced forms of SARP and SEMP model that contain spatial lags of independent variables with order 1 to infinity only differ in their coefficient terms. On this basis, they argue that reduced forms of the SEMP and SARP cannot, in practice, be empirically distinguished from a specification with spatial lags of exogenous variables, if the real data generating process is unknown. The reason for this result is that some assumption has to be made regarding the form of the weights matrix $\mathbf{W_T}$ and the spatial lags of different orders on the explaining variables are almost always expected to be highly correlated. Invoking the Gibbons and Overman logic we decided to estimate a further specification that complements the previous models with spatial lags of the exogenous regressors (SLXP) on the base on Pooled OLS and Fixed Effects Estimators, namely:

$$\mathbf{E} = \gamma_0 \boldsymbol{\iota}_{\boldsymbol{nT}} + \mathbf{Y_N}\boldsymbol{\gamma}_{\mathbf{t}} + \theta_0 \mathbf{P} + (\boldsymbol{\theta}_{\mathbf{t}}^{\mathbf{IDID}\prime}\mathbf{Y_N})'\mathbf{P} + \mathbf{X}\boldsymbol{\delta} + \mathbf{W}_{\boldsymbol{T}}\boldsymbol{P}\boldsymbol{\vartheta} + \boldsymbol{\varepsilon} \tag{12}$$

The spatial lag term $\mathbf{W}_{\boldsymbol{T}}\boldsymbol{P}\boldsymbol{\vartheta}$ includes spatial lags of the regional Kaitz index.

Interim results

We present estimates of the DID model (1) and its spatial counterparts using (the log of) employment as the labour market outcome of interest to summarise the NMW effect on employment over the medium term, namely the average over 14 years since its introduction relative to the base period of 1997/98. We do this for each of our three geographies 140, 138 and 406. These are presented in columns 1 of Tables 1, 2 and 3 respectively. Since our estimates are identified by variations in the NMW bite over time across areas, the coefficients of the Kaitz NMW toughness measure suggest that there is no overall difference in employment growth rates between areas where the NMW bites most compared to areas where the NMW has less impact. Column 7 to 12 of each of those tables include the GVA lagged variable as a measure for the recession. In each case this recession proxy is always positively significant suggesting that employment is always higher when there is higher economic growth (as measured by GVA lagged.) These estimates are reported as a simple benchmark for our more sophisticated models.

Columns 1 and 6 of Tables 4, 5 and 6 augment the base model with the model specification of the IDiD estimator in equation (5) for each geography (140, 138 and 406 respectively) where we do not model the spatial nature of the data. The second 6 columns in these Tables also include the GVA lagged variable as a measure for the recession. As expected, the recessionary variable is always positively significant suggesting the intuitively 'correct' sign of the impact of growth on employment. It should be noted that the addition of the GVA variable always attenuates downwards the size of the IDiD coefficients on the NMW variables for each year. Even in this benchmark model this indicates that modeling the employment effect of the MW without taking account of demand shocks and recessions is problematic and likely to overstate any measured MW effects on employment.

Columns 2, 3, 7 and 8 of Tables 1 to 3 present the results of the limited model in equation (1) but in the spatial context using the SEMP model from equation (11) excluding the Incremental Diff-in-Diff term $(\boldsymbol{\theta}_{\mathbf{t}}^{\mathbf{IDID}\prime}\mathbf{Y}_{\mathbf{N}})'\mathbf{P}$. The next set of estimates for each geography is presented in columns 2, 3, 7 and 8 of Tables 4 to 6 which estimates the 'full model' using the complete Incremental Diff-in-Diff structure but includes the spatial effects using the SEMP model from equation (11) for respectively the 140, 138 and 406 samples. The main results from our estimations are suggested by the patterns in these tables. First taking the common findings across nearly all specifications we can suggest that the recession, as captured by the GVA lagged variable, plays an important role in the determination of employment but that the consequence of this variable's inclusion is that the NMW interaction effects are always attenuated. Likewise these estimated effects are further attenuated when we explicitly take account of the spatial effects.

Our second main finding is that in our specifications with area effects the coefficient on the Kaitz index does not have a significant effect on employment in the the presence of the NMW . After including the IDiD Kaitz interaction term this overall effect gets nearly in all specifications significantly negative whereas the coefficient of the Kaitz interaction term becomes predominantly positively significant. This can be interpreted as the continuous, underlying effect of having a NMW in place rather than the effect of the size of the year on year up rating. This is an important conclusion which potentially enables us to understand much of the controversy in the research literature. Indeed it suggests that, if our specification is correct (and our logic were applicable to the US state context), then the source of much of the disagreement between the main protagonists, Card and Krueger (1994, 1995) or Neumark and Wascher (1994, 2004) may have been misplaced due to equation misspecification.

Turning to the estimations and distinguishing between the results from the different geographies is instructive. Looking first at Tables 1 and 4 relating to the 140 geography we see that there are clearly positively significant interaction effects in the years 2003 to 2007 inclusive. Comparing these tables to Tables 2 to 5 relating to the 406 geography we see that the size of the effects is attenuated but that nearly the same years are positively significant at the 10% level and very nearly significant at the 5% level. This is a consistent finding which directly relates to the extra variance introduced by having such a finer geographical set of areas with more variance in outcomes, more potential for measurement error and unobserved heterogeneity.

The more unusual findings are those which relate to the 138 areas which are travel to work areas (TTWA) in Tables 3 to 7. Here we see in Table 5 that there is no evidence of interaction effects in the IDiD model with the exception of the years 2003 and 2004. This suggests that if one uses a geography which is defined on the nature of the commuting pattern - which the TTWA is - then this effectively cancels out the IDiD effect. Hence, if one believes that the analysis should be done on the basis of the TTWA geography, then there are no appreciably significant incremental year on year effects of the NMW.

Turning to the models which reflect the specification of the spatial models (columns 3 to 6 and 9 to 12) we found for the SEMP specifications a significantly positive coefficient λ for the spatial lag of the error terms for the model versions without the recession variable and in the case of the 406 geography for all versions of the SEMP. These ancillary parameters nearly always become insignificant whenever the GVA lagged variable is included in the 140 geography. The explanation is that the GVA variable is measured at the Government Office Regional level which is the higher administrative geographical unit to the 140 geography and, in effect, the spatial dimension is soaked up by the inclusion of this variable. This is reflected in the fact that the lack of significance of λ tells us that the spatial model is not necessary (when the GVA lagged term is included). The results of Lagrange Multiplier tests, presented in the next section, confirm that with the exception of the full specified models.

Overall, where our NMW incremental effects are found significant it should be stressed that these point estimates effects are small in magnitude, but it is clear that they are masked if the simple DiD Policy-On Policy -Off variable is used. If the standard assumptions of Diff-in-Diff relating to the Stable Unit Treatment are applicable (namely that no other systematic factors are varying across geography and over time) then we can interpret this as a direct impact of the year on year up-ratings to the NMW which may cancel out the overall negative impact of the presence of the NMW as captured by the Kaitz variable. On this basis, if anything, employment rate appears to have risen more in areas where the NMW has more relevance

[Insert Table 1 Here]

[Insert Table 2 Here]

[Insert Table 3 Here]

[Insert Table 4 Here]

[Insert Table 5 Here]

[Insert Table 6 Here]

Dynamics and endogenous regressors.

A standard assumption of the OLS and Fixed effects models as well as their spatial counterparts, (including the SEMP, SARP and SLXP models described and estimated in the previous sections), is that they require that the explanatory variables are uncorrelated with the residuals. In practical applications like ours, this requirement is rarely completely satisfied. Potential reasons for this could be the dynamic properties of the used variables, e.g. hysteresis of the employment variable, measurement errors in the variables or further omitted variables that are not observable[15]. To overcome these problems the most commonly used approach is to use dynamic panel instrumental variable methods (see Arellano and Bond 1991, Greene 2012, pp. 256 f.f.).[16]

We adopt the system generalized method of moments estimator (SGMM) developed by Holtz-Eakin, Newey and Rosen (1988), Arellano and Bover (1995), and Blundell and Bond (1998). Generally, the SGMM estimator can be applied for panel data sets with a short observation period in terms of small T and many cross section units, thus large N (Roodman 2009a). Furthermore, the estimator assumes that the only available instruments are "internal" in terms of lags of the instrumented variables in differences or levels. We prefer the SGMM to Difference GMM (and other alternatives) because SGMM is more efficient and precise as it reduces the finite sample bias (Baltagi 2008). Nevertheless, a crucial initial condition for the validity of the additional instruments in SGMM is that fixed effects don't correlate with differences in the instrumenting variables. Roodman (2009b) showed that this requirement can be fulfilled in a "steady state" situation, when deviations from long-term values are not systematically related to fixed effects. A "steady state" can be assumed, if the variable of interest – here it is the employment rate – tends to converge. This can be checked empirically by unit root tests for panel models or by estimating a simple fixed effects

[16] The estimation strategy behind the assumption that some or all explaining variables are correlated with the error term is that one finds a set of (instrument) variables that are correlated with the explaining variables but not with the error terms. Due to the resulting set of relationships among instruments, explaining variables, and error terms a consistent estimator of the coefficients of interest can be derived. For this purpose a number of assumptions have to be made, Roodman (2009a).

model with only a lagged dependent variable as regressor (AR(1) model). In the latter case the coefficient of the lagged dependent variable should be smaller than the absolute value of unity. The results of these estimations confirm that our lagged dependence coefficients are significantly smaller than unity[17] for all our alternative geographical areas.

Our baseline equations are nearly equivalent to the fully specified SLXP models from equation (12) with and without the recession variable:

$$E_{jt} = \alpha E_{jt-1} + \gamma_0 + J_j + \sum_{i=1999}^{2010} \gamma_i Y_i + \theta_0 P_{jt} + \sum_{k=1999}^{t} \theta_k^{IDID} Y_k\, P_{jk} + \rho \sum_{i=1}^{n} w_{ij} P_{it} + \pi D_{jt} + \delta X_{jt} + \varepsilon_{jt} \tag{13}$$

We included the lagged employment rate E_{jt-1} and complemented the model with certain levels and differences of the control variables (in X_{jt}) as instruments. In what follows we will refer to that model as SGMM-SLXP model.[18] Furthermore, in our baseline specification we excluded the direct Kaitz effect coefficient $\theta_0 P_{jt}$ for two reasons: (1) the Kaitz index measured in the observed region is often suspected to be endogenous; (2) since the Kaitz index highly correlates with the spatial lag of the Kaitz coefficient[19], thus the (weighted) average of the Kaitz index in related regions, we use the SLXP term $\sum_{i=1}^{n} w_{ij} P_{it}$ as instrument and interpret its coefficient ρ as the nearest parameter for the true direct effect of the MW.

In our estimation strategy an important consideration is to find the optimal specification in terms of the choice and number of the instruments. This is important since too many instruments could 'over-fit' the endogenous variables (of interest) and lead to invalid results (Roodman 2009b). Regarding the choice of the instruments, we already gave reasons why the overall MW bite, as well as the yearly incremental MW bites, are strictly exogenous and not influenced by their lagged level or difference values. In contrast, since we have to bear in mind that we apply a reduced form of employment equation since we cannot observe the detailed employment generating process at the micro-level, it could suggest that the employment rate[20], as well as the control variables, are also

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17 We estimated the equation $y_{it} = \alpha y_{it-1} + \mu_i + \epsilon_{it}$ with the employment rate y_{it}, y_{it-1} in region *i* and time *t* and its value from the pre-period *t-1*, the lagged dependent coefficient α, the fixed effect μ_i, and the error term ϵ_{it}. The results are (1) $\alpha_{140\,areas} =$ 0.234*** , (2) $\alpha_{138\,areas} =$ 0.199***, and (3) $\alpha_{406\,areas} =$ 0.184***. All coefficients are significantly smaller than unity (significance level of 1 per cent). Details and results of the unit root tests can be provided upon request.

18 The properties of the SGMM estimator are described in the literature . The SLXP term does not change the relevance and exclusion restriction for the instruments (Gibbons and Overman, 2012).

19 See also the details in Appendix F.

20 Why didnt we include a lagged term of the dependent variable in the OLS, FE, and spatial specifications we present here. The reason is that it has to be expected that pooled OLS specification deliver an upward bias of the appropriate coefficient and FE downward biased results. However, beside others we used adequate specifications to detect the validity of the SGMM. See also below in the text.

partly influenced by their previous values. In all our estimations, we use the two step procedure to get robust results and test statistics as well as the Windmeijer correction for small sample size to reduce the downward bias of standard errors after the two step procedure (Windmeijer 2005). Furthermore to handle instrument proliferation we integrated the instrument set into one column as it is proposed by Roodman (2009a). We utilze the Arellano-Bond test for AR(1) and AR(2) in first differences, Hansen's *J*-test of overidentifying restrictions, and the Difference-in-Hansen's *J* statistic to test the validity of the model specification[21]. Furthermore, we checked the robustness of the model specification by varying the number of lags of the instrumental variable sets (Roodman 2009b). In line with Bond (2002) we checked if the coefficient for the lagged dependent variable lies somewhere between the results obtained from adequate OLS and FE estimators.[22] Throughout these specification tests, our reported results provide acceptable test values ensuring that our estimates remain robust.

Table 7 contains the results. All other result tables are in Appendix D. The first two columns in Table 7 contain the estimation results for the 140 areas, the second two columns for the 138 areas and the last two columns for the 406 areas. Though we found slightly different results depending on the geography used, there is considerable concordance with the patterns of the estimates we found in our baseline estimations. An expected and common result is the significant positive coefficient of the lagged dependent employment variable.

[Insert Table 7 Here]

Regarding the employment effects of the NMW the coefficient of the spatial lag of the Kaitz indices based on the commuting matrix is significantly negative whereas the same coefficient based on contiguity – though it has a negative sign in all versions – is insignificant. The coefficient for the demand variable has a positive sign, nevertheless it does not significantly differ from zero in all model versions of the 140 and 406 areas.

The results for the estimation of the coefficients of the Incremental-Differences-in-Differences interaction terms differ somewhat by the three geographies. For the 140 regions we found partly positive effects on a significance level of at least 10 per cent for the years 2004, 2006 and 2007. For the 138 regions there are partly significantly positive effects in 2009 and there is one specification based on the contiguity matrix with even a negative employment effect in 2002 (column 6). The

[21] A description and the results of the mentioned tests can be found in Appendix F.

[22] Though we don't report the results from Pooled OLS and FE specifications including lagged dependent variables, the results can be provided by the authors upon request.

results for the geography with 406 areas reveal significant positive coefficients for the year 2000 and again a negative effect for one specification based on contiguity in 1999 (column 12).

We can now summarize the conclusions of our empirical work in three figures which compares the nature of our estimates over our different specifications. This is provided in Figures 6 to 8 for the 140, 138, and 406 areas. These figures include 6 separate panels representing the key specifications we have estimated. Starting from the top left panel we see the FE estimates (those results reported in column. 1 of Table 4 to 6). Here we see 8 of the interaction coefficients are significant. This is the result reported in the Dolton et al (2010) paper which suggest the separate identification of year on year interaction effects. The second panel reports what happens to these parameters when the GVA demand shock variable is included – we see that none of the interaction terms are now significant. The third panel reports the SLXP model – where we see that also 8 of these interactions are significant; and in the fourth panel – the SLXP model with the GVA variable – the interaction terms are again insignificant. The fifth and the sixth panel show the SGMM-SLXP specification either on the base of the commuting or the contiguity matrix. In the commuting version only two interaction terms are statistically significant and in the contiguity version there is no interaction effect significantly different from zero. The last two specifications report the effect of modelling spatial dependence, shocks in demand, the endogeneity of the Kaitz, and the autoregressive nature of the steady state employment process. Our conclusion is that the effect of the interaction terms modeling the year on year uprating disappear when a more rigorous approach to the key econometric problems of modelling the employment effect of the NMW is adopted.

These results suggest that the modeling of demand shocks and the pattern of the employment is important for the identification of year on year effects. Quite clearly these effects are severely attenuated by more rigorous econometric models and estimation methods. Reassuringly our results are robust to the fairly stern test of using different geographical units of observation and also to using a completely different weights matrix to model the spatial dependence.[23]

[Insert Figure 5 Here]

[Insert Figure 6 Here]

[Insert Figure 7 Here]

Our empirical strategy is strongly substantiated by examination the residuals for spatial autocorrelation. We found weak spatial autocorrelation in our pure fixed effects model specifications, but after including the demand variable spatial autocorrelation cannot be neglected.

[23] We also checked the robustness of or results using restricted geography samples by dropping out those regions that are economically weak. Though we use regional variation for our identification strategy, we assume that the economic behavior in terms of the employment elasticity of the bite of the minimum wage is equal over all regions. It is known that the economic power of the UK is not equally distributed over the whole country and there are some economic hot spots like – first of all –London, but also the whole area around London and there are other regions that are more or less economically dependent of those strong regions. Hence, it could be reasonable that the employment effects of the MW are somewhat different between the strong and the weaker regions. Details can be found in Appendix E.

Therefore it could be important the demand variable refers to a higher aggregation level than our observed geographies. To test for spatial autocorrelation we use a Moran's I test as well as Pesaran's CD test.[24] The advantage of Moran's I is that it can be utilized to test for spatial autocorrelation in a given spatial dependence structure before and after a model is applied. So it could be prescriptive in seeking an adequate model specification. Nevertheless, at the same time this a shortcoming when it is generally of interest that the residuals are not spatially correlated. Another crucial shortcoming is that Moran's *I* is quite sensitive to outliers in the spatially related data as well as in the weights matrix, this can easily lead to spurious conclusions. Furthermore, it is obviously that the results of Moran's *I* can also be sensitive to the choice of the form of the weights matrix. Pesaran (2004 argues that in panel data it is preferable to test the residuals without an *apriori* assumption about the structure of the spatial dependence. The same author proposes a simple test statistic, named Pesaran's CD test (Pesaran 2004, 2010, 2011). The statistic is reported to be robust to structural breaks and adequate to panel data where the number of cross sections is larger than the number of observations periods as it is here the case for all three geographies. Alhough, we also report the results for Moran's *I* as well as for Pesaran's *CD* test in Appendix F for illustration purposes, we summarize here the results of the latter test. These test statistics imply the presence of spatial dependence in the residuals after including the demand variable (GVA). This can possibly be explained by the fact that the GVA variable relates to larger spatial aggregates than the regions in all geographies. With only a few exceptions it can be shown that the models, that consider spatial dependencies accordingly, reduce the extent of the spatial dependency in the residuals. The results also provide the preferred SGMM-SLXP models including only the SLXP term $\sum_{i=1}^{n} w_{ij} P_{it}$ without a direct Kaitz measure. The coefficient of those variable is insignificant in all model variations and including it seems only to distort the residuals towards a spatial pattern in our specifications.

To sum up, our findings suggest that the most reliable results can be derived from the SGMM-SLXP model with a modified Kaitz measure based on the commuting matrix since this matrix represents an appropriate approximation to the asymmetric spatial dependence structure which characterizes UK regional labour markets. Following the results of the spatial autocorrelation tests we further conclude that the problem of spatial autocorrelation in the residuals is best overcome with the 406 geography regions. The results imply a small significant negative long-term employment effect. The elasticity is around of around .1 implying that a 10 per cent increase in the bite of the minimum wage (in terms of the Kaitz index) would lead to a fall of 1.1 per cent of the

[24] Details for both tests can be found in Appendix H.

employment rate. Our results also strongly suggest that there is no incremental year-on-year effect of the uprating of the Minimum Wage.

4 Conclusions

The contribution of this paper is to bring up to date the econometric evidence of the impact of the NMW on employment in the UK and focus on the particular context of the recession of 2008-10. We use geographical variation in the impact of the NMW and the recession to identify the separate employment consequences of imposing a NMW and its up-ratings over the years.

We used four sources of variation to try and identify the effect of the NMW in the UK. The first is to exploit a natural variation in how the NMW bites in different geographical locations. In our UK case the MW is set nationally, thus there is no decision to be made at the local level (in sharp contrast to the US State case). This implies a natural variation in the exact bite of the NMW which is different at each geographical area. Our second source of variation was to examine the effect of changes in the up-rating of the NMW over the years since it was introduced. This estimation is based on an Incremental Diff-in-Diff method which allows us to estimate the marginal (interaction) effect of each year's change in the NMW. The third source of variation we exploit is to allow the size, timing and duration of the recession to affect different regions differently. This provides us with the capacity to estimate the effect of the recession on local employment and to net out this factor in assessing the impact of the NMW on employment. Clearly, any assessment which fails to net out this factor will bias the estimates of the impact of the NMW on employment. The final source of variation we exploit is that our spatial model allows the nature of the complex pattern of regional interconnectedness and spillover effects to be present in the estimation and thus gave us some confidence that we have netted out their impact in terms of the effect of the NMW on employment. The combination of these four different sources of variation in the data along with the rigorous use of different robustness checks means that we can be more confident about the estimated effect of the impact of the NMW. Our conclusions are all the more credible in the sense that we got largely consistent results even though we reanalyzed the data using three completely different geographies.

Our first concern is that all the existing literature on the MW is unable to distinguish between having a statutory MW in existence and the changes or up-ratings of the MW, e.g. on an annual basis. We are able to make such a distinction since we are able to observe a 'pre-period' before the MW legislation was enacted (namely prior 1999), when no NMW existed. It turns out that such an important distinction is vital to try and understand why the previous literature has found negative or zero employment effects of the MW.

Our second major concern was directed at the literature which has attempted to use geographical variation in the bite of the MW to identify its effects. The problem with this literature is that it has ignored the possible geographical contiguity and spatial spill overs associated with neighbouring locations. More specifically the literature to date has modeled panel data on geographical areas over time under the assumption that they are completely independent observations. We relaxed this assumption using the latest spatial econometric methods to explore the effect of location spill-over effects on the relationship between the MW and employment outcomes. To do this we used the observed pattern of commuting behaviour between our geographical areas as this represents the reduced form pattern of the many observed and unobserved influences on commuting decisions and hence the interconnectedness of local labour markets. For robustness we compare our results with model specifications which utilize contiguity matrices instead and found that our results were largely unchanged.

Most of the literature on the employment impact of the MW has ignored the potential identification problem associated with netting out the effect of changes in the aggregate economy. The third contribution of this paper is to condition directly on the nature of demand shocks in our estimation of the regional employment effects of the MW. This is most pertinent since we are interested in the effect the MW in a time of the deepest recession the western world has experienced since the Great Depression of the 1930s. We attempted to solve this problem by using geographically varying information from Gross Value Added by region which is related to the onset, severity and duration of the recession in different locations. We find that this important adaptation of the standard econometric model leads directly to a considerable attenuation of the year on year interaction effects and hence explains why previous papers got (spuriously) positive employment effects.

Our fourth area of contribution relates directly to the modeling of the dynamic form of employment changes. Specifically we use a SGMM model to circumvent potential endogeneity problems relating to the inclusion of a lagged term in the employment dependent variable. This modeling adopts an Arellano-Bond (1991) type IV estimation procedure to overcome the bias introduced by having a dynamic model of employment adjustment. We found that when the effects of demand side shocks, a lagged dependent variable and endogenous wage effects are explicitly modeled then the effects on the NMW on employment that have been previously found are largely attenuated. Specifically we find that the incremental effects on NMW upratings are not identified but that the effect of having a NMW is predominately small and negative. It is quite clear that the modeling of changing demand conditions is a real contribution of this paper which suggests that firstly – papers in the past which have neglected to study this may be finding spuriously significant effects when none exist and secondly, that the overall effect of NMW depends on what stage of the cycle the economy is in. We can see that a tight labour market in a recession may induce negative employment effects – but that NMW uprating in boom times may induce little or no overall effect. This relationship points the importance of controlling adequately for demand side factors (including the steady state autoregressive path of employment) in modeling what might happen to employment in the face of a NMW (imposition or up-rating). Our further robustness checks suggest that our preferred models do not exhibit spatial dependence.

Our final area of contribution was to tackle the thorny issue of the endogeneity of the Kaitz variable in the model. Clearly the concern is that the relative measure of the MW variable may be directly affected by the level of employment or unemployment in the economy and hence the simultaneous use of this variable as if it were an exogenous explanatory variable may be spurious. To circumvent this problem we adopted a novel spatial IV which used the value of the Kaitz index in the neighbouring areas as an instrument for the Kaitz index in any given geography on the grounds that this weighted measure should be correlated with the level of the Kaitz in the area but potentially uncorrelated with the specific unobserved heterogeneity in the area. Our results suggested that this was a useful solution to the problem of endogeneity and revealed rigorous conclusions.

The conclusion from our estimates is that overall there are no incremental employment effects of up ratings to the MW in a year on year context. The years where the more naive estimation strategies revealed a small positive coefficient are 2003, 2004, 2005 and 2006 which are historically the years when the NMW up rating exceeded the RPI rise in the cost of living and hence the up rating of the NMW was relatively generous and where there is a boom in the economy and hence a potential measurement error problem in the modeling of employment. Allowing for our more robust estimation procedures shows that, in contrast, the underlying effect of the presence of the NMW is reflected in the Kaitz Index coefficient in the tables. This coefficient is nearly always negative and significant suggesting that the effective implementation of the NMW has an underlying negative impact on employment. It should be stressed that our measured marginal effects were consistently attenuated when we condition out for the presence and severity of the recession in the regional context. Our conclusions were robust to different definitions of the geography used to perform the estimation. Furthermore, they are robust to changing the nature of our weights matrix. Specifically we used a simple neighbouring contiguity matrix instead of our commuting matrix and our overall conclusions did not change.

Our findings are interesting as they rationalize the controversial debate in the literature as to why one might get negative impacts of the MW – i.e. due to the effect of the presence of the MW rather than its up rating. Our results are also consistent with much of the recent literature focusing on the introduction of the NMW but also because they explain why it may be possible to get both zero and positive effects. Our results thus present quite a radical departure from the literature which has studied the employment effects of the minimum wage but never distinguished between the effect of imposing a MW and uprating the MW on a regular basis. We can suggest that the overall effect of having introduced the MW could have small but clear negative employment consequences. In contrast we can see that the effect of uprating a MW will nearly always be insignificantly different from zero. We have also demonstrated that the reason for some of literature finding positive effects of the MW is that it does not distinguish between the issues of: spatial dependence, the endogeneity of the MW (in the form of the Kaitz index), recessionary demand shocks, and the steady state trend in the employment series. Our suggestion from this UK evidence is that failing to take account of these complications could lead to spuriously positive MW effects with underestimated standard errors. Although our evidence is only for one county there are grounds to support the hypothesis that our results can provide a methodology which can reconcile the perennial debate between the pro and anti-MW lobbies.

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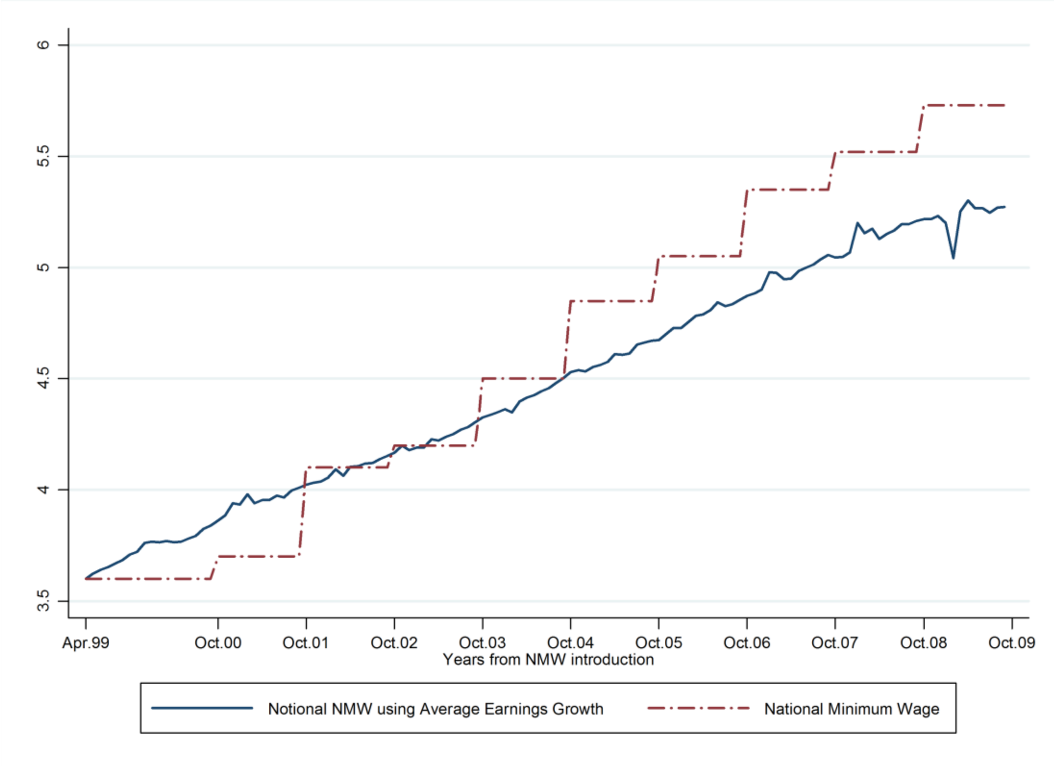


Figure 1. Change in Nominal Hourly Wage Level of the NMW

Source: NMW: UK Low Pay Commission, www.lowpay.gov.uk; Average Earnings Growth: http://data.gov.uk/dataset/average_earnings_index

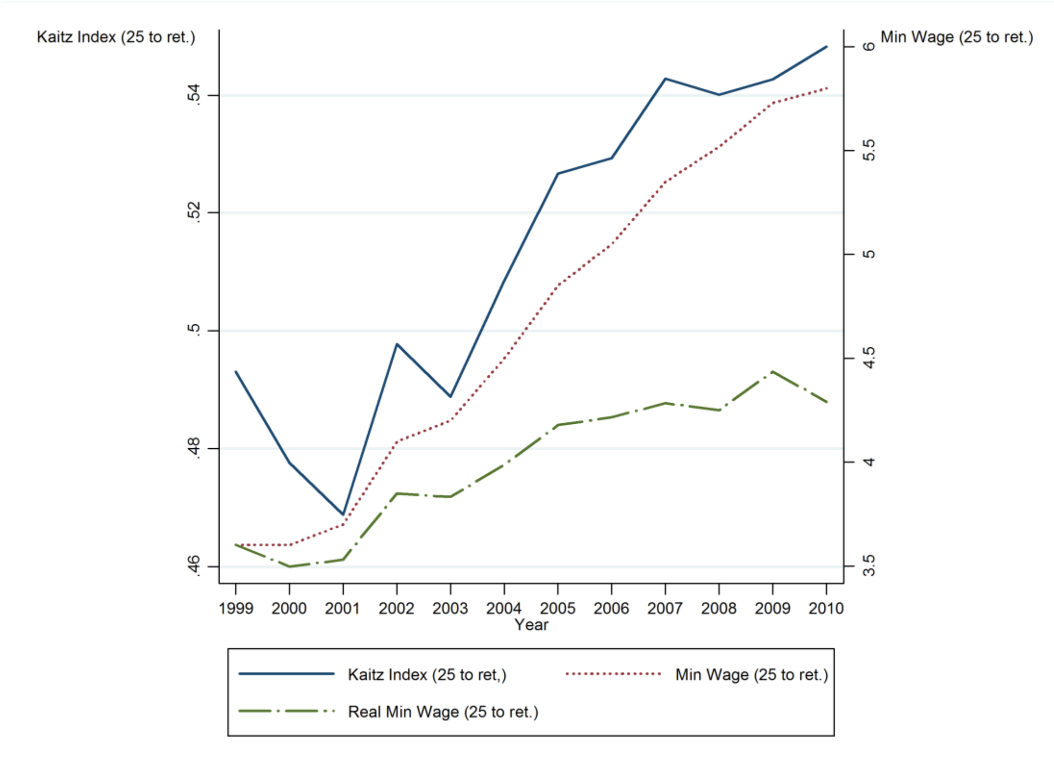


Figure 2. Change in the Estimated NMW and Kaitz Index over Time

Source: NMW: UK Low Pay Commission, www.lowpay.gov.uk; Kaitz Index: ASHE, own calculation; Real NMW: deflated by RPI, http://www.ons.gov.uk/ons/publications/re-reference-tables.html?edition=tcm%3A77-260874

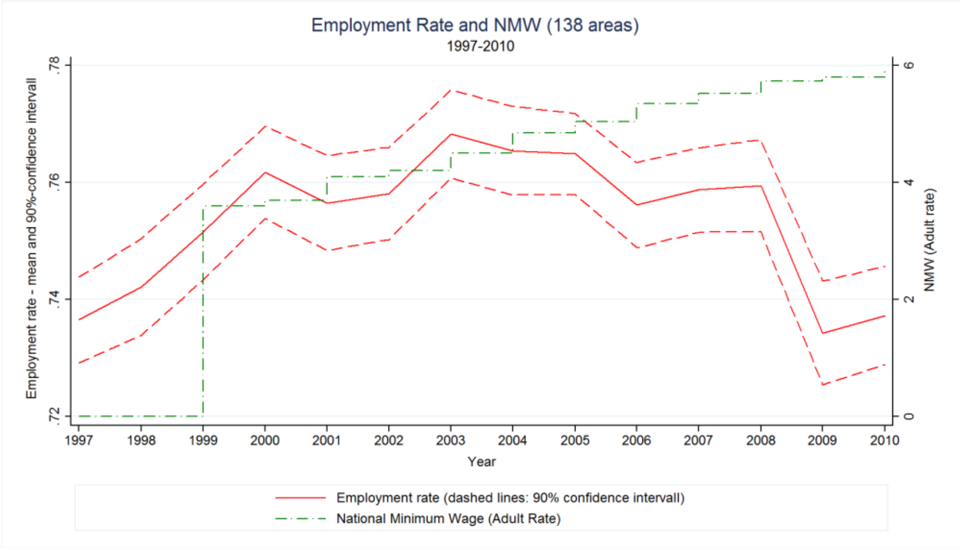


Figure 3. National Minimum Wage (Adult Rate) and employment rate in the UK 1997-2010.

Source: NMW: UK Low Pay Commision, www.lowpay.gov.uk; unweighted yearly averages of the Employment Rate, own calculation.

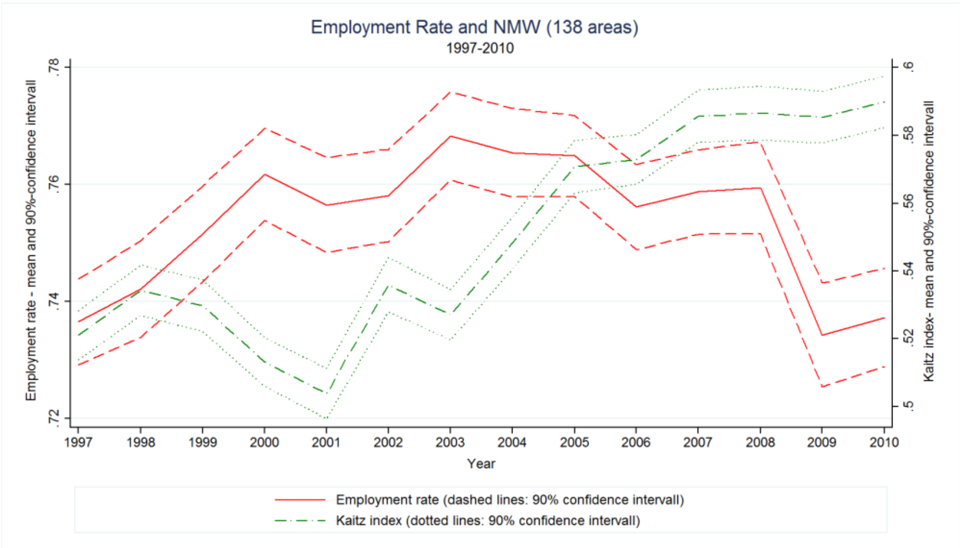


Figure 4. Kaitz index (Adult Rate) and employment rate in the UK 1997-2010.

Note: unweighted yearly averages of the Employment Rate and the Kaitz index (MW/Median), own calculation.

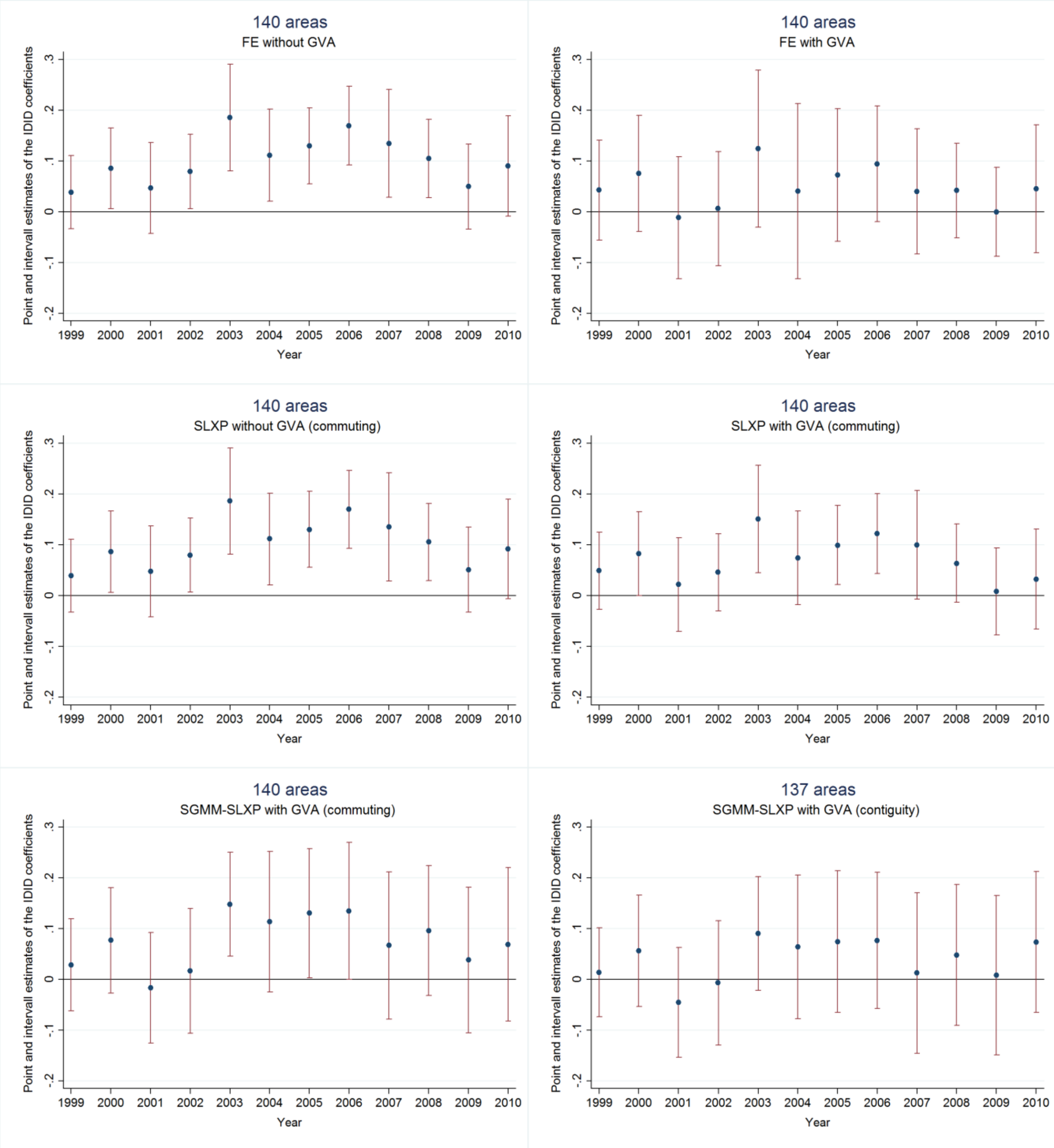


Figure 5. Results of the Point and Intervall Estimations of the Incremental-Difference-in-Difference Coefficients for the SLXP and the SGMM-SLX models for 140 Areas.

Note: 95% confidence intervals; SGMM-SLXP models without direct Kaitz effect. In SLXP Models with contiguity matrix the islands are excluded.

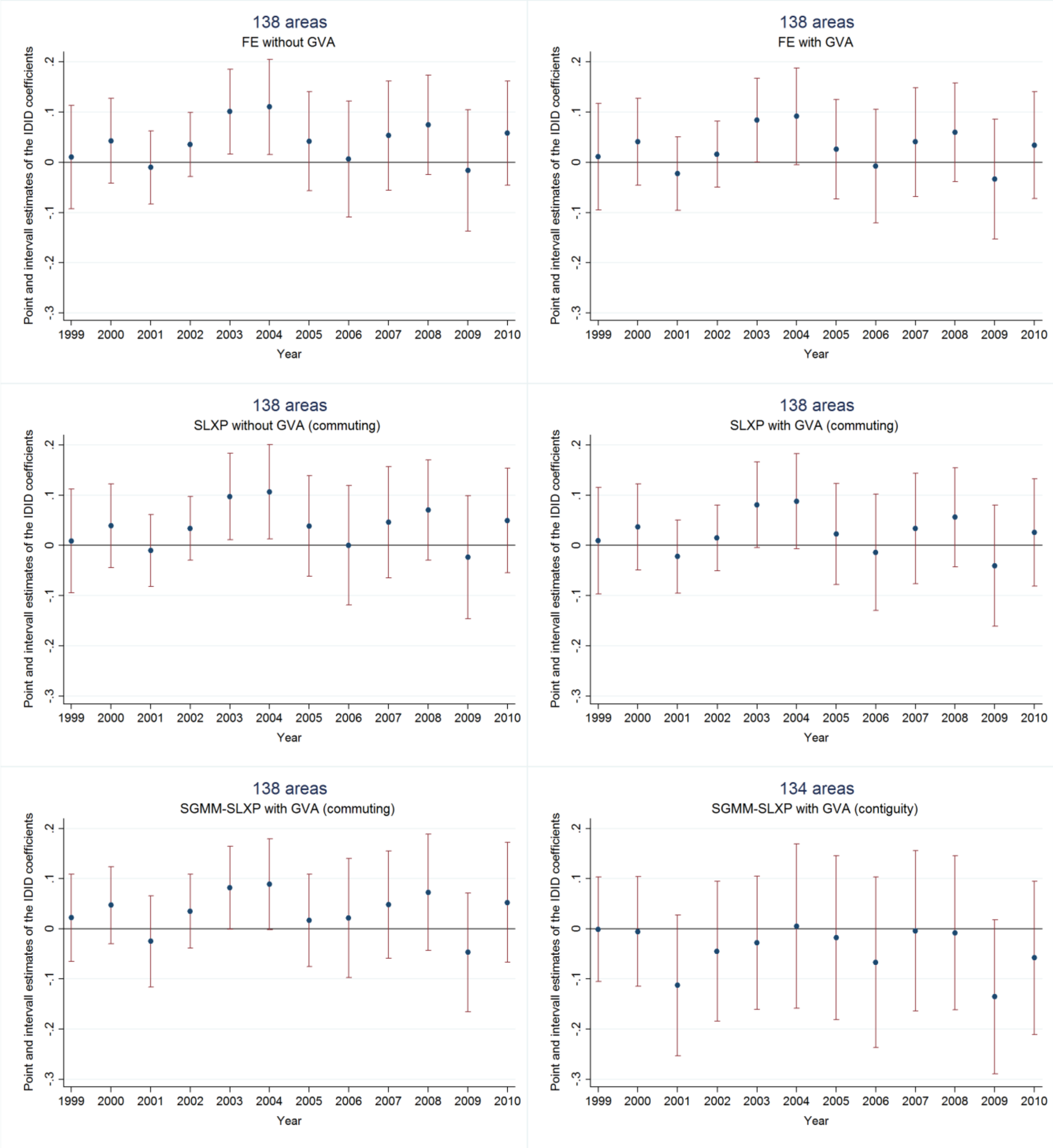


Figure 6. Results of the Point and Intervall Estimations of the Incremental-Difference-in-Difference Coefficients for the SLXP and the SGMM-SLX models for 138 Areas.

Note: 95% confidence intervals; SGMM-SLXP models without direct Kaitz effect. In SLXP Models with contiguity matrix the islands are excluded.

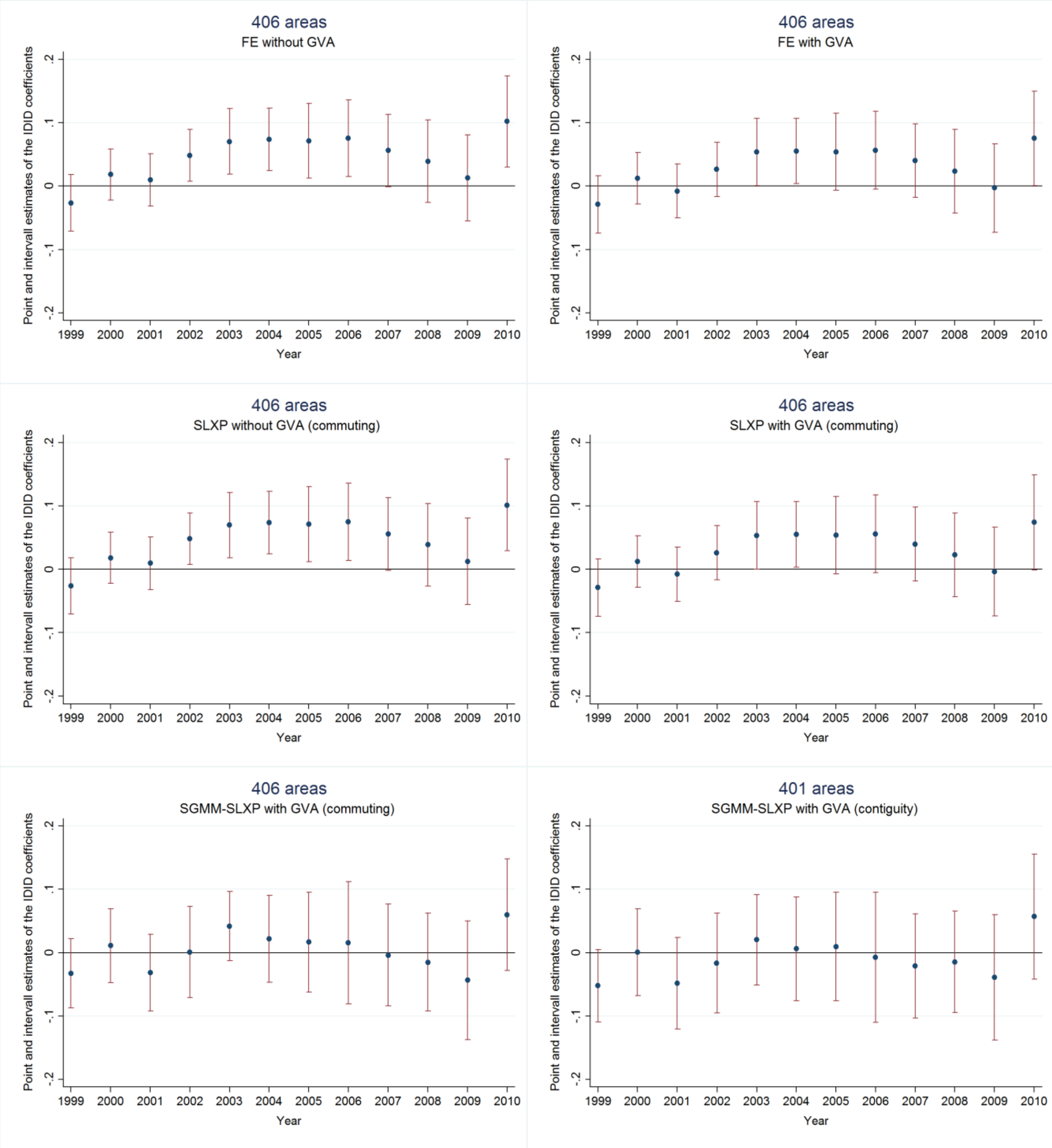


Figure 7. Results of the Point and Intervall Estimations of the Incremental-Difference-in-Difference Coefficients for the SLXP and the SGMM-SLX models for 406 Areas.

Note: 95% confidence intervals; SGMM-SLXP models without direct Kaitz effect. In SLXP Models with contiguity matrix the islands are excluded.

Table 1. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 140 areas level, 1997-2010, all regressions contain control variables, area and year effects.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix
Kaitz Index	-0.056	-0.054*	-0.055*	-0.058	-0.048	-0.049	-0.049	-0.049	-0.053	-0.047
	(0.044)	(0.032)	(0.032)	(0.043)	(0.044)	(0.044)	(0.032)	(0.032)	(0.044)	(0.045)
w*Kaitz Index				0.016	-0.065				0.042	-0.021
				(0.094)	(0.066)				(0.091)	(0.063)
GVA						0.966***	0.958***	0.968***	0.971***	0.916***
						(0.205)	(0.160)	(0.161)	(0.205)	(0.207)
Share Public	-0.076*	-0.074**	-0.074**	-0.076*	-0.069*	-0.071*	-0.071**	-0.071**	-0.072*	-0.066*
	(0.040)	(0.034)	(0.034)	(0.040)	(0.040)	(0.039)	(0.033)	(0.033)	(0.039)	(0.039)
age	-3.949	-3.806	-3.865*	-3.938	-4.796	-3.782	-3.855*	-3.826*	-3.754	-4.566
	(2.651)	(2.314)	(2.320)*	(2.656)	(2.958)	(2.578)	(2.298)	(2.298)	(2.582)	(2.854)
age2	0.098	0.095	0.096*	0.098	0.120	0.094	0.096*	0.095*	0.093	0.114
	(0.066)	(0.058)	(0.058)	(0.067)	(0.074)	(0.065)	(0.058)	(0.058)	(0.065)	(0.072)
age3	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001*	-0.001	-0.001	-0.001
	(0.001)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.001)	(0.001)
nvq4plusIMP	0.179***	0.171***	0.174***	0.179***	0.177***	0.164***	0.163***	0.163***	0.165***	0.164***
	(0.040)	(0.034)	(0.034)	(0.040)	(0.041)	(0.041)	(0.034)	(0.034)	(0.040)	(0.042)
total_female	0.072	0.066	0.071	0.072	0.071	0.043	0.043	0.043	0.043	0.043
	(0.065)	(0.054)	(0.054)	(0.065)	(0.066)	(0.063)	(0.054)	(0.054)	(0.063)	(0.065)
lambda		0.088**	0.045				0.016	0.018		
		(0.037)	(0.031)				(0.038)	(0.031)		
Observations	1,960	1,960	1,960	1,960	1,918	1,960	1,960	1,960	1,960	1,918
R-squared	0.115			0.115	0.118	0.132			0.133	0.134
ll		3,334.000	3,332.588				3,351.555	3,351.591		

Notes: (Robust) standard errors in parentheses (columns 1,4,5,6,9,10)
*** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 2. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 138 areas level, 1997-2010, all regressions contain control variables, area and year effects.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix
Kaitz Index	-0.108***	-0.099***	-0.103***	-0.099**	-0.091**	-0.102**	-0.096***	-0.098***	-0.094**	-0.089**
	(0.039)	(0.035)	(0.035)	(0.039)	(0.043)	(0.039)	(0.035)	(0.035)	(0.039)	(0.043)
w*Kaitz Index				-0.120	-0.070				-0.112	-0.062
				(0.119)	(0.086)				(0.117)	(0.084)
GVA						0.742***	0.701***	0.749***	0.736***	0.591***
						(0.226)	(0.192)	(0.188)	(0.226)	(0.225)
Share Public	-0.062	-0.062*	-0.063*	-0.060	-0.072*	-0.068*	-0.067**	-0.068**	-0.066	-0.077*
	(0.040)	(0.033)	(0.033)	(0.040)	(0.041)	(0.040)	(0.033)	(0.033)	(0.041)	(0.042)
age	2.486	2.28	2.553	2.538	1.253	2.105	1.989	2.162	2.156	1.267
	(2.455)	(1.807)	(1.808)	(2.440)	(3.216)	(2.534)	(1.804)	(1.804)	(2.520)	(3.266)
age2	-0.060	-0.055	-0.062	-0.062	-0.029	-0.051	-0.048	-0.053	-0.052	-0.029
	(0.061)	(0.045)	(0.045)	(0.060)	(0.080)	(0.063)	(0.045)	(0.045)	(0.062)	(0.081)
age3	0.000	0.000	0.001	0.001	0.000	0.000	0.000	0.000	0.000	0.000
	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.000)	(0.000)	(0.001)	(0.001)
nvq4plusIMP	0.159***	0.155***	0.156***	0.158***	0.150***	0.149***	0.148***	0.147***	0.148***	0.142***
	(0.035)	(0.028)	(0.028)	(0.035)	(0.038)	(0.034)	(0.028)	(0.028)	(0.035)	(0.037)
total_female	0.034	0.028	0.029	0.031	0.068	0.023	0.02	0.021	0.021	0.057
	(0.066)	(0.050)	(0.050)	(0.066)	(0.065)	(0.066)	(0.050)	(0.050)	(0.065)	(0.065)
lambda		0.136***	0.080**				0.105***	0.066**		
		(0.038)	(0.033)				(0.039)	(0.033)		
Observations	1,932	1,932	1932	1,932	1,876	1,932	1932	1932	1,932	1,876
R-squared	0.131			0.132	0.133	0.139			0.140	0.138
ll		3,240.368	3,238.184				3,247.243	3,246.62		

Notes: (Robust) standard errors in parentheses (columns 1,4,5,6,9,10)
*** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 3. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 406 areas level, 1997-2010, all regressions contain control variables, area and year effects.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix	FE	SEMP Commuting Matrix	SEMP Contiguity Matrix	SLXP Commuting Matrix	SLXP Contiguity Matrix
Kaitz Index	-0.030	-0.028	-0.029	-0.028	-0.024	-0.030	-0.028	-0.029	-0.028	-0.025
	(0.023)	(0.019)	(0.019)	(0.024)	(0.024)	(0.023)	(0.019)	(0.019)	(0.024)	(0.024)
w*Kaitz Index				-0.036	-0.050				-0.035	-0.044
				(0.053)	(0.042)				(0.053)	(0.042)
GVA						0.638***	0.598***	0.633***	0.637***	0.580***
						(0.147)	(0.129)	(0.124)	(0.146)	(0.148)
Share Public	0.004	0.002	0.003	0.004	0.003	-0.001	-0.001	-0.001	-0.001	-0.001
	(0.027)	(0.022)	(0.022)	(0.027)	(0.027)	(0.026)	(0.022)	(0.022)	(0.026)	(0.027)
age	1.417	1.479	1.464	1.400	1.224	1.514	1.547*	1.54*	1.497	1.372
	(1.273)	(0.903)	(0.905)	(1.274)	(1.322)	(1.277)	(0.901)	(0.903)	(1.278)	(1.331)
age2	-0.036	-0.037*	-0.037	-0.035	-0.031	-0.038	-0.039*	-0.039*	-0.038	-0.035
	(0.032)	(0.023)	(0.023)	(0.032)	(0.033)	(0.032)	(0.022)	(0.023)	(0.032)	(0.033)
age3	0.000	0.000*	0.000*	0.000	0.000	0.000	0.000*	0.000*	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
nvq4plusIMP	0.181***	0.178***	0.179***	0.180***	0.178***	0.177***	0.175***	0.176***	0.176***	0.175***
	(0.024)	(0.017)	(0.017)	(0.024)	(0.025)	(0.024)	(0.017)	(0.017)	(0.024)	(0.025)
total_female	0.007	0.004	0.006	0.006	0.009	0.005	0.002	0.004	0.004	0.007
	(0.038)	(0.030)	(0.030)	(0.038)	(0.038)	(0.037)	(0.030)	(0.030)	(0.037)	(0.037)
lambda		0.112***	0.048**				0.094***	0.039*		
		(0.024)	(0.020)				(0.024)	(0.020)		
Observations	5,684	5,684	5,684	5,684	5,614	5,684	5,684	5,684	5,684	5,614
R-squared	0.064			0.065	0.066	0.069			0.070	0.070
ll		8,053.414	8,046.582				8,064.743	8,060.443		

Notes: (Robust) standard errors in parentheses (columns 1,4,5,6,9,10); *** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 4. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 140 areas level, 1997-2010, all regressions contain control variables, area and year effects..

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity
Kaitz Index	-0.137***	-0.136***	-0.137***	-0.140***	-0.133***	-0.106**	-0.106***	-0.106***	-0.110**	-0.108**
	(0.050)	(0.041)	(0.041)	(0.050)	(0.049)	(0.051)	(0.041)	(0.041)	(0.052)	(0.051)
w*Kaitz Index				0.027	-0.058				0.042	-0.024
				(0.091)	(0.064)				(0.090)	(0.063)
GVA						0.888***	0.890***	0.886***	0.891***	0.807***
						(0.221)	(0.168)	(0.168)	(0.220)	(0.224)
Share Public	-0.072*	-0.072**	-0.071**	-0.072*	-0.064	-0.073*	-0.073**	-0.073**	-0.073*	-0.066
	(0.040)	(0.034)	(0.034)	(0.040)	(0.041)	(0.040)	(0.033)	(0.033)	(0.040)	(0.041)
Kaitz*1999	0.038	0.041	0.039	0.039	0.039	0.048	0.048	0.048	0.049	0.050
	(0.036)	(0.045)	(0.044)	(0.036)	(0.037)	(0.039)	(0.043)	(0.043)	(0.038)	(0.039)
Kaitz*2000	0.085**	0.088**	0.086*	0.086**	0.090**	0.081*	0.08*	0.081*	0.082*	0.086**
	(0.040)	(0.045)	(0.044)	(0.041)	(0.040)	(0.042)	(0.043)	(0.044)	(0.042)	(0.042)
Kaitz*2001	0.047	0.048	0.047	0.047	0.050	0.021	0.02	0.021	0.022	0.026
	(0.045)	(0.043)	(0.042	(0.045)	(0.046)	(0.047)	(0.042)	(0.042)	(0.047)	(0.047)
Kaitz*2002	0.079**	0.079*	0.07*	0.079**	0.085**	0.046	0.045	0.046	0.046	0.054
	(0.037)	(0.043)	(0.042	(0.037)	(0.037)	(0.038)	(0.042)	(0.042)	(0.038)	(0.039)
Kaitz*2003	0.185***	0.184***	0.18***	0.186***	0.185***	0.150***	0.15***	0.15***	0.151***	0.152***
	(0.053)	(0.044)	(0.044	(0.053)	(0.055)	(0.054)	(0.043)	(0.044)	(0.054)	(0.056)
Kaitz*2004	0.111**	0.108**	0.1**	0.112**	0.114**	0.074	0.074*	0.074*	0.074	0.078
	(0.046)	(0.045)	(0.044	(0.046)	(0.047)	(0.047)	(0.044)	(0.044)	(0.047)	(0.048)
Kaitz*2005	0.130***	0.128***	0.12***	0.130***	0.137***	0.099**	0.098**	0.099	0.099**	0.109***
	(0.038)	(0.047)	(0.046)	(0.038)	(0.038)	(0.039)	(0.046)	(0.046)	(0.039)	(0.039)
Kaitz*2006	0.169***	0.166***	0.16***	0.170***	0.173***	0.121***	0.122***	0.122***	0.122***	0.128***
	(0.039)	(0.047)	(0.046	(0.039)	(0.040)	(0.040)	(0.046)	(0.047)	(0.040)	(0.041)
Kaitz*2007	0.135**	0.132***	0.13***	0.135**	0.137**	0.099*	0.099**	0.099**	0.100*	0.103*
	(0.054)	(0.046)	(0.046	(0.054)	(0.055)	(0.054)	(0.046)	(0.046)	(0.054)	(0.055)
Kaitz*2008	0.105***	0.101**	0.10**	0.105***	0.111***	0.063	0.063	0.063	0.063	0.070*
	(0.039)	(0.045)	(0.04)	(0.039)	(0.040)	(0.039)	(0.045)	(0.045)	(0.039)	(0.041)
Kaitz*2009	0.049	0.047	0.049	0.051	0.075*	0.006	0.006	0.006	0.008	0.034
	(0.042)	(0.047)	(0.046)	(0.042)	(0.040)	(0.043)	(0.046)	(0.046)	(0.043)	(0.042)
Kaitz*2010	0.090*	0.087*	0.089*	0.092*	0.099**	0.030	0.03	0.03	0.032	0.044
	(0.050)	(0.047)	(0.047)	(0.050)	(0.049)	(0.050)	(0.048)	(0.048)	(0.050)	(0.050)
lambda		0.037	0.01				-0.01	-0.003		
		(0.038)	(0.031)				(0.038)	(0.031)		
Observations	1,960	1,960	1,960	1,960	1,918	1,960	1,960	1,960	1,960	1,918
R-squared	0.130			0.130	0.133	0.143			0.143	0.144
ll		3,348.821	3,348.417				3,363.179	3,363.164		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10). *** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 5. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 138 areas level, 1997-2010, all regressions contain control variables, area and year effects..

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity
Kaitz Index	-0.145***	-0.134***	-0.139***	-0.134***	-0.130**	-0.127***	-0.122***	-0.123***	-0.117**	-0.119**
	(0.047)	(0.044)	(0.044)	(0.047)	(0.051)	(0.048)	(0.044)	(0.044)	(0.048)	(0.052)
w*Kaitz Index				-0.107	-0.057				-0.106	-0.054
				(0.120)	(0.086)				(0.119)	(0.085)
GVA						0.709***	0.678***	0.719***	0.708***	0.545**
						(0.239)	(0.192)	(0.190)	(0.238)	(0.241)
Share Public	-0.059	-0.060*	-0.060*	-0.058	-0.070*	-0.066	-0.065**	-0.066**	-0.064	-0.075*
	(0.040)	(0.033)	(0.033)	(0.040)	(0.042)	(0.040)	(0.033)	(0.033)	(0.041)	(0.042)
Kaitz*1999	0.010	0.011	0.012	0.009	0.009	0.011	0.012	0.012	0.010	0.011
	(0.052)	(0.051)	(0.050)	(0.052)	(0.052)	(0.054)	(0.050)	(0.049)	(0.054)	(0.054)
Kaitz*2000	0.043	0.045	0.041	0.040	0.046	0.041	0.043	0.039	0.037	0.045
	(0.043)	(0.051)	(0.050)	(0.042)	(0.043)	(0.044)	(0.050)	(0.049)	(0.043)	(0.044)
Kaitz*2001	-0.010	-0.009	-0.009	-0.010	-0.013	-0.023	-0.02	-0.021	-0.022	-0.022
	(0.037)	(0.049)	(0.048)	(0.036)	(0.036)	(0.037)	(0.049)	(0.048)	(0.037)	(0.036)
Kaitz*2002	0.035	0.03	0.032	0.034	0.038	0.016	0.015	0.014	0.015	0.023
	(0.032)	(0.049)	(0.048)	(0.032)	(0.032)	(0.033)	(0.048)	(0.048)	(0.033)	(0.033)
Kaitz*2003	0.101**	0.093*	0.096*	0.098**	0.095**	0.084**	0.08	0.08	0.081*	0.082*
	(0.043)	(0.050)	(0.049)	(0.044)	(0.043)	(0.042)	(0.050)	(0.049)	(0.043)	(0.042)
Kaitz*2004	0.110**	0.105**	0.110**	0.107**	0.109**	0.091*	0.090*	0.092*	0.088*	0.094*
	(0.048)	(0.050)	(0.049)	(0.048)	(0.047)	(0.048)	(0.050)	(0.049)	(0.048)	(0.048)
Kaitz*2005	0.042	0.032	0.038	0.039	0.043	0.026	0.021	0.024	0.023	0.031
	(0.050)	(0.052)	(0.051)	(0.051)	(0.052)	(0.050)	(0.052)	(0.051)	(0.051)	(0.052)
Kaitz*2006	0.006	-0.006	-0.003	0.000	0.007	-0.008	-0.015	-0.015	-0.014	-0.005
	(0.058)	(0.054)	(0.053)	(0.060)	(0.060)	(0.057)	(0.053)	(0.052)	(0.059)	(0.060)
Kaitz*2007	0.053	0.046	0.045	0.047	0.044	0.040	0.037	0.034	0.034	0.033
	(0.055)	(0.053)	(0.052)	(0.056)	(0.055)	(0.055)	(0.052)	(0.051)	(0.056)	(0.055)
Kaitz*2008	0.074	0.072	0.075	0.071	0.070	0.060	0.06	0.061	0.056	0.058
	(0.050)	(0.051)	(0.050)	(0.050)	(0.051)	(0.050)	(0.051)	(0.050)	(0.050)	(0.052)
Kaitz*2009	-0.016	-0.018	-0.017	-0.023	0.006	-0.034	-0.032	-0.032	-0.040	-0.009
	(0.061)	(0.053)	(0.051)	(0.062)	(0.068)	(0.060)	(0.052)	(0.051)	(0.061)	(0.067)
Kaitz*2010	0.058	0.052	0.052	0.050	0.070	0.034	0.034	0.031	0.026	0.050
	(0.053)	(0.053)	(0.052)	(0.053)	(0.052)	(0.054)	(0.053)	(0.052)	(0.054)	(0.054)
lambda		0.122***	0.068**				0.09**	0.061*		
		(0.039)	(0.033)				(0.039)	(0.033)		
Observations	1,932	1,932	1,932	1,932	1,876	1,932	1,932	1,932	1,932	1,876
R-squared	0.137			0.138	0.139	0.145			0.145	0.144
ll		3,246.423	3,244.811				3,252.833	3,252.477		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10). *** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 6. Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 406 areas level, 1997-2010, all regressions contain control variables, area and year effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity	FE	SEMP Commuting	SEMP Contiguity	SLXP Commuting	SLXP Contiguity
Kaitz Index	-0.065**	-0.06**	-0.063***	-0.063**	-0.060**	-0.051*	-0.049**	-0.05**	-0.049*	-0.049*
	(0.028)	(0.025)	(0.024)	(0.028)	(0.028)	(0.028)	(0.025)	(0.025)	(0.028)	(0.029)
w*Kaitz Index				-0.015	-0.038				-0.019	-0.036
				(0.053)	(0.042)				(0.053)	(0.042)
GVA						0.497***	0.47***	0.497***	0.498***	0.429***
						(0.160)	(0.134)	(0.129)	(0.160)	(0.161)
Share Public	0.005	0.004	0.004	0.005	0.004	0.001	0.001	0.001	0.002	0.001
	(0.026)	(0.022)	(0.022)	(0.026)	(0.027)	(0.026)	(0.022)	(0.022)	(0.026)	(0.027)
Kaitz*1999	-0.026	-0.027	-0.027	-0.026	-0.026	-0.029	-0.029	-0.029	-0.029	-0.028
	(0.023)	(0.027)	(0.027)	(0.023)	(0.023)	(0.023)	(0.027)	(0.027)	(0.023)	(0.023)
Kaitz*2000	0.018	0.019	0.018	0.018	0.022	0.012	0.014	0.012	0.012	0.017
	(0.021)	(0.027)	(0.026)	(0.021)	(0.021)	(0.021)	(0.027)	(0.026)	(0.021)	(0.021)
Kaitz*2001	0.010	0.008	0.009	0.010	0.010	-0.008	-0.007	-0.008	-0.008	-0.005
	(0.021)	(0.026)	(0.026)	(0.021)	(0.021)	(0.022)	(0.026)	(0.026)	(0.022)	(0.022)
Kaitz*2002	0.049**	0.046*	0.047*	0.049**	0.051**	0.026	0.027	0.025	0.026	0.032
	(0.021)	(0.026)	(0.026)	(0.021)	(0.021)	(0.022)	(0.026)	(0.026)	(0.022)	(0.022)
Kaitz*2003	0.070***	0.066**	0.069***	0.070***	0.068**	0.054**	0.052*	0.053**	0.053*	0.054**
	(0.026)	(0.027)	(0.026)	(0.026)	(0.026)	(0.027)	(0.027)	(0.026)	(0.027)	(0.027)
Kaitz*2004	0.074***	0.067**	0.071***	0.074***	0.074***	0.055**	0.052*	0.054**	0.055**	0.058**
	(0.025)	(0.027)	(0.026)	(0.025)	(0.025)	(0.026)	(0.027)	(0.027)	(0.026)	(0.026)
Kaitz*2005	0.071**	0.066**	0.07**	0.071**	0.072**	0.054*	0.051*	0.053*	0.054*	0.057*
	(0.030)	(0.028)	(0.028)	(0.030)	(0.030)	(0.031)	(0.028)	(0.028)	(0.031)	(0.031)
Kaitz*2006	0.075**	0.067**	0.073***	0.075**	0.075**	0.057*	0.052*	0.055*	0.056*	0.059*
	(0.031)	(0.029)	(0.028)	(0.031)	(0.031)	(0.031)	(0.029)	(0.028)	(0.031)	(0.031)
Kaitz*2007	0.056*	0.049*	0.053*	0.056*	0.054*	0.041	0.036	0.039	0.040	0.041
	(0.029)	(0.028)	(0.027)	(0.029)	(0.029)	(0.029)	(0.028)	(0.027)	(0.030)	(0.030)
Kaitz*2008	0.039	0.029	0.037	0.039	0.038	0.023	0.017	0.022	0.023	0.024
	(0.033)	(0.027)	(0.027)	(0.033)	(0.033)	(0.034)	(0.028)	(0.027)	(0.034)	(0.034)
Kaitz*2009	0.013	0.009	0.012	0.012	0.016	-0.003	-0.004	-0.003	-0.004	0.002
	(0.035)	(0.028)	(0.028)	(0.035)	(0.035)	(0.036)	(0.028)	(0.028)	(0.036)	(0.036)
Kaitz*2010	0.102***	0.097***	0.100***	0.101***	0.102***	0.075**	0.074**	0.074**	0.074*	0.079**
	(0.037)	(0.029)	(0.028)	(0.037)	(0.037)	(0.038)	(0.029)	(0.029)	(0.038)	(0.038)
lambda		0.094***	0.034*				0.080***	0.026		
		(0.024)	(0.020)				(0.024)	(0.020)		
Observations	5,684	5,684	5,684	5,684	5,614	5,684	5,684	5,684	5,684	5,614
R-squared	0.071			0.071	0.072	0.074			0.074	0.074
ll		8,070.099	8,065.145				8,076.612	8,073.021		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10). *** p<0.01, ** p<0.05, * p<0.1; In SLXP Models with contiguity matrix the islands are excluded.

Table 7. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, excluding direct effect of Minimum Wage, 16 years to retirement age, 1997-2010, all regressions contain control variables, area and year effects..

	(1)	(2)	(3)	(4)	(5)	(6)
	140 regions		138 regions		406 regions	
	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Commut.
Kaitz Index	-	-	-	-	-	-
	-	-	-	-	-	-
w*Kaitz Index	-0.183***	-0.174***	-0.260***	-0.204***	-0.107**	-0.105**
	(0.059)	(0.060)	(0.063)	(0.049)	(0.042)	(0.041)
E_{t-1}	0.254***	0.247***	0.237***	0.238***	0.227***	0.227***
	(0.032)	(0.031)	(0.042)	(0.035)	(0.024)	(0.024)
GVA		0.243		0.411		0.142
		(0.270)		(0.315)		(0.190)
Share Public	-0.137***	-0.140***	-0.074	-0.100**	-0.180***	-0.181***
	(0.041)	(0.041)	(0.049)	(0.041)	(0.030)	(0.030)
Kaitz*1999	0.012	0.028	0.008	0.085*	-0.033	-0.033
	(0.049)	(0.046)	(0.066)	(0.048)	(0.028)	(0.028)
Kaitz*2000	0.059	0.077	-0.048	0.072*	0.013	0.011
	(0.059)	(0.052)	(0.066)	(0.043)	(0.030)	(0.030)
Kaitz*2001	-0.031	-0.017	-0.128**	-0.021	-0.027	-0.031
	(0.060)	(0.055)	(0.058)	(0.052)	(0.030)	(0.031)
Kaitz*2002	0.004	0.016	-0.061	0.037	0.007	0.001
	(0.066)	(0.062)	(0.058)	(0.053)	(0.035)	(0.037)
Kaitz*2003	0.140**	0.148***	0.003	0.073	0.044	0.042
	(0.054)	(0.052)	(0.056)	(0.053)	(0.027)	(0.028)
Kaitz*2004	0.102	0.113	-0.042	0.115**	0.025	0.021
	(0.073)	(0.070)	(0.079)	(0.055)	(0.034)	(0.035)
Kaitz*2005	0.115	0.130**	-0.086	0.054	0.021	0.016
	(0.070)	(0.064)	(0.067)	(0.064)	(0.039)	(0.040)
Kaitz*2006	0.125*	0.135*	-0.091	0.024	0.021	0.015
	(0.070)	(0.068)	(0.070)	(0.065)	(0.048)	(0.049)
Kaitz*2007	0.052	0.067	-0.093	0.115*	0.000	-0.004
	(0.078)	(0.073)	(0.087)	(0.064)	(0.040)	(0.041)
Kaitz*2008	0.083	0.096	0.017	0.097	-0.011	-0.015
	(0.069)	(0.065)	(0.069)	(0.067)	(0.039)	(0.039)
Kaitz*2009	0.020	0.038	-0.097	-0.063	-0.038	-0.044
	(0.079)	(0.073)	(0.072)	(0.074)	(0.047)	(0.048)
Kaitz*2010	0.056	0.069	-0.035	0.061	0.065	0.060
	(0.080)	(0.077)	(0.082)	(0.070)	(0.044)	(0.045)
Observations	1,820	1,820	1,794	1,794	5,278	5,278
Number of instruments	69	70	69	70	69	70

*** p<0.01, ** p<0.05, * p<0.1. Test statistics are provided in Table 8.

Table 8. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, test statistics of estimates in table 7

	(1)	(2)	(3)	(4)	(5)	(6)
	140 regions		138 regions		406 regions	
	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA
Number of instruments	69	70	69	70	69	70
	Arellano-Bond test for AR in first differences					
AR(1)	-6.3686	-6.3565	-6.4099	-6.3256	-13.9387	-13.9078
Prob > z	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
AR(2)	1.0287	1.0033	-0.5672	-1.0680	-0.3281	-0.3141
Prob > z	0.3036	0.3157	0.5706	0.2855	0.7429	0.7534
	Hansen test of overidentified restrictions					
J	28.2410	28.9745	43.7693	41.9121	40.2927	39.7507
Prob> chi2	0.8185	0.7907	0.1751	0.2298	0.2860	0.3066
	Difference-in-Sargan tests of exogeneity of instrument subsets					
	- GMM instruments -					
C	26.2418	26.7218	43.0096	40.9723	36.1964	35.7007
Prob > chi2	0.7919	0.7717	0.1139	0.1605	0.3217	0.3426
	- Instrumented variables in levels and first differences -					
C	16.9532	16.8779	23.6900	22.0738	25.3033	22.0680
Prob > chi2	0.5930	0.5315	0.2083	0.2287	0.1508	0.2290
C	18.6466	18.6739	26.5089	27.7266	25.6241	25.1688
Prob > chi2	0.7706	0.7691	0.3279	0.2717	0.3725	0.3966

5 Appendix

A Data

Geography of 140, 406 and 138 areas

In this paper we use three different levels of geographical aggregation. The first two geographies (WLA and WAREA) are borne of administrative areas. The last level of aggregation approximates travel-to-work areas (TTWAs).The WLA geography includes 32 London boroughs, 238 Districts[25], 36 metropolitan districts and the 46 Unitary Authorities in England; the 22 Unitary Authorities in Wales and 32 Unitary Councils in Scotland, resulting in 406 (WLA) local areas. The median ASHE sample cell size is 291 and the smallest cell is 36. Figure A-1 shows what this geography looks like. The WAREA geography includes 34 English counties, 6 English metropolitan counties, 46 English Unitary authorities, Inner and Outer London and finally 52 Unitary authorities in Scotland and Wales. This results in 140 local areas[26], shown in Figure A-2. Here the median sample cell size is 581 and the smallest cell is 53. These two administrative geographies are related between each other. Unitary authorities and Scottish Unitary councils are both present in the WLA and in the WAREA levels of aggregation. What differentiate the two geographies is that the small English districts which characterize the WLAs are aggregated into Counties in the WAREAs. Also the 32 London boroughs of the WLAs are combined into Inner and Outer London in the WAREAs. Figure A-3 clearly shows how different London looks for these two levels of aggregation.

Our last level of aggregation aims to approximate travel to work areas (TTWAs) which, following an ONS definition, correspond to areas in which 67% of people living and working in the same geography. Since TTWAs are not available for the entire period of analysis the only option was to attempt to replicate our results for the most reasonable definition of TTWA that we could manually reconstruct from the data available. We use ASHE data from 2002 where we have information about the WLAs where people work and WLAs where people live. We then compute commuting shares (given by the proportion of people who live in an area and work in another area and the proportion of people who work in an area and live in another one). We then keep all the districts and unitary authorities where the ONS definition of travel to work areas holds (around 12% of areas). For the other WLAs, with the help of GIS software we overlap the map of ONS TTWA with the map of WLAs and combine Districts and Unitary Authorities into existing TTWA boundaries. With these new geographies we compute the commuting patterns to check the consistency with ONS definition of

[25] The London borough City of London and the district Isles of Scilly are excluded from the analysis due to small sample sizes.

[26] The Orkney Islands, Shetland Isles and Western Isles are aggregated together. The 36 English metropolitan districts are combined into 6 English Metropolitan Counties. London Boroughs are aggregated into Inner and Outer London. This allows to have match geographies in the LFS and in the ASHE/NES, using the definition of the variable "uacnty" in the LFS.

travel-to-work areas. For the few areas (14%) where the ONS definition of travel-to-work areas still does not hold we aggregate further. Some 90% of these are such that at least 67% of working residents work in the area and at least 67% of workers are resident in the areas. This gives us 138 (new geographical) TTWA areas, showed in Figure A-4, for which we repeated our analysis. Here the median sample cell size is 581 and the smallest cell is 37. [27]

Figure A-5 helps to understand how our TTWAs actually differ from our WLAs and WAREAs. By focusing only on a small part of the country such as the south-London coast, we can see how the TTWAs are generally extending over the narrow boundaries of the Districts, which characterize the WLAs. They can also be smaller than the administrative counties of the WAREAs. Also the administrative boundaries of the counties do not necessarily determine the boundaries of the TTWA, since people living at the borders of a county can commute and work in a neighboring town which is not necessarily part of that specific county. This figures clearly shows the merits and limitations of working with different units of geography. Focus on the county for Kent (the county with the darker shading) - the most southeastern of the counties in the UK - which is at the bottom left hand corner of our figures in Fig A-5 in a) b) and c). Using the WAREA geography we see that this whole area is basically a single geography with the sole exception of Gillingham and Chatham which adds additional further urban area in the north west of the county. Using TTWA the county becomes 4 different areas based on the feasibility of the transport connections (and by definition the observed patterns of commuting behavior). At the same time it loses a small slice of its southwestern edge to neighbouring Sussex. Using the WLA geography the county becomes 10 separate areas which more evenly divide the county on approximately equal square surface areas. So clearly the 406 areas will combine the detail of the old administrative units with an element of the logistic issue of being local areas of a small enough size to facilitate commuting behavior. It will also have the advantage, from the spatial geographic perspective that each geography is likely to include its own centroid. This is less likely with the larger geographical areas. Whilst it is not necessarily the case that this logic for Kent applies to the whole country the pattern we have described here is nevertheless indicative of what is true for the rest of the country. Based on this logic we would favour the 406 as our unit of analysis for this empirical work. Whilst this gives us the advantage of having a much larger dataset (and concomitant degrees of freedom for our econometric model), the obvious limitation of using this geography is that the variables we derive are based on smaller underlying micro dataset sample sizes (and thus have potentially larger measurement error).

Definition of key variables

Employment rate

Total number of employees, self-employed, unpaid family workers and participants in government-supported training and employment programs in working age as a proportion of people in working age in each local area.

Data on employment used in this paper is taken from June to August of each year.

Source: Labour Force Survey. Residence based analysis.

Kaitz Index

The ratio of the NMW to the median hourly wage in each local area:

- Starting from 1999, the shares are a weighted average of the minimum wage shares of persons from 18 to 21 years and of persons from 22 to retirement age.
- From 2004, with the introduction of the new development rate for young between 16 and 17 years, the shares are a weighted average of the minimum wage shares of persons of persons of 16 and 17 years, of persons from 18 to 21 years and of persons from 22 to retirement age.

Generally the ASHE/NES based minimum wage variables used in this paper are recorded in April of each year and the NMW variables are recorded six months after each NMW up-rating due to the fact that the minimum wage was invoked in April 1999 but then up-rated in October of each following year. There are however two exceptions: April 1999 which is contemporaneous to the introduction of the minimum and April 2000, which is one year from the introduction of the minimum. To reduce simultaneity concerns, the wage data in April of year t is regarded as having absorbed any effect of the NMW upgrade in October of year $t-1$. This is in turn matched to employment data taken from June to August of year t. For the pre-period 1997 and 1998, data on employment rates are collected from March 1997 to February 1998 and from March 1998 to February 1999. Quarterly data is not available for these two calendar years. Since LFS Local Area data is only available in seasonal quarters, it is only possible to use the quarter June-August and not a longer period (eg. from May to September). This means that the estimated impact effect we identify is a mixture of the impact of the up-rating in year $t-1$ and any changes from the already announced anticipation of the effect of the new NMW level in year t. As a robustness check we have varied our timing assumptions and our results suggest that any anticipation effect is negligible. Swaffield (2008) shows that there is very little early upward adjustment in wages in the six months prior to October over several years of data.

Other Covariates

The other covariates in the dataset are derived from the underlying micro datasets. In the case of age, gender and sector we use the ASHE/NES to compute these variables by each of the geographies and simply computed proportions. In the case of the human capital regressor – the proportion of the workforce with a degree level qualification we used the LFS to derive this variable since the ASHE/NES does not have education information on respondents. The Gross Value Added (GVA) variable was derived from official statistics sources from Regional Trends from various years Table 3.3 Workplace-based gross value added1 (GVA) per head at current basic prices. This GVA variable is measured at the level of the government office region and not at the level of the individual geography. It is not possible to measure the effect of aggregate demand at any finer geography than the government office region level. However this has the advantage that we are controlling for demand shocks at a different level of geography than our basic units of observation.

Further properties of the ASHE/NES datasets to be considered

Even if ASHE is considered to give reliable wage figures though payroll records and it has a relatively large sample size, there are some limitations of this dataset which could affect this study.

- Possible measures of hourly earnings

  The Low Pay Commission recommended construction of the hourly pay variable on the NES/ASHE data involves dividing gross pay (excluding overtime, shift and premium payments) by basic paid hours. This variable closely matches the definition of National Minimum Wage. However, the variable is available in the panel only from 2000. It is therefore necessary to use another measure of hourly earnings in this study which covers the period 1997 to 2007.

  The variable used is a "basic hourly wage rate", defined as gross weekly earnings excluding overtime, and divided by normal basic hours. As a result this variable will be slightly larger than the true hourly wage and the measurement error will tend to be larger, the higher shift and premium payments are. This might therefore result in an under-statement of the number of low paid workers.

- Discontinuities in NES/ASHE dataset across years

  Time series analysis has been complicated when the ASHE replaced the NES in 2004 and also by several changes in the ASHE methodology from 2004 to 2007.

  First of all, the coverage of employees for the ASHE is greater than that of the NES. The NES surveys employees taken from HM Revenue & Customs PAYE record, excluding the majority of those whose weekly earnings fall below the PAYE deduction threshold. Moreover, this survey does not cover employees between sample selection for a particular year and the survey reference week in April. Thus, mobile workers who have changed or started new jobs between the drawing of the sample and the reference week are excluded. In conclusion, NES understate the proportion on NMW as it does not record the earnings of many low paid workers, especially part-time and mobile workers. In 2004, ASHE survey was introduced to improve on the representation of the low paid: it improved coverage of employees including mobile workers who have either changed or started new jobs between sample selection and the survey reference in April. Also, the sample was enlarged by including some of the employees outside the PAYE system.

In 2005 a new questionnaire was introduced. In particular, the definition of incentive/bonus pay changed to only include payments that were paid and earned in April. Also, a new question including "pay for other reasons" was introduced. This implies respondents might include earnings information which was not collected in the past. Even if results for 2004 have been reworked to exclude irregular bonus/incentive payments and to allow for this missing pay, results from 1997 to 2003 remain inconsistent with the ones from 2004 onwards.

Given that the main source of information on hourly pay in this study includes shift and premium payments and from 2004 "pay for other reasons", estimations of measures of minimum wage and wage inequality might be affected by this discontinuity, with an increase of the average measurement error and the dispersion in the measurement error from 2004 onwards.

Finally, in 2007 the sample size of ASHE was reduced by 20%. ASHE results for 2007 are based on approximately 142,000 returns, down from 175,000 in 2006. The largest sample cuts occurred principally in industries where earnings are least variable, affecting the randomness of the sample.

Consistent series which takes into account of the identified changes has been produced going back from 2006 to 2004 and from 2007 to 2006. For 2004 results are also available that exclude supplementary information, to be comparable with the back series generated by imputation and weighting of the 1997 to 2003 NES data. Unfortunately, it is not possible to get consistent datasets for the entire period concerning this study (1997-2007).

B Weights matrix in spatial specifications

We compare our results with model specifications containing contiguity matrices. Hereby we decided to handle the islands as regions without neighbours. That means that the rows and columns of the contiguity matrix for the islands (Isle of Anglesey, Isle of Wight, Orkney Islands, Shetland Islands, and Western Isles) only contains Zeros and consequently they are assumed not to interact with other regions. Compared to the results for the community matrix we found weaker spatial interdependency coefficients. We explain that with the fact that commuting represents the spatial interdependencies of local labour markets in a more satisfactory way than simply weighting by which other areas each location shares a geographical border. However, it is important that we examine the extent to which our results are robust to the assumption that the commuting matrix is not the sole form of weights matrix which will give rise to these results. Hence, we use the contiguity matrix to establish this. We found that the results for the NMW coefficients are robust to the variation of the weights matrix.

C Lagrange multiplier tests for spatial specification

We used the Lagrange Multiplier tests described and provided by Debarsy and Ertur (2010) for the panel models with area effects to shed more light on these questions from a statistical point of view. Therefore at first we test for the null hypothesis of (i) a joint spatial lag of the dependent variables and spatial autoregressive error terms vs. there is at least one kind of spatial dependency. In this and all other tests the Null hypothesis has to be rejected if the p value is higher than a certain significance level. Afterwards if the null hypothesis has to be rejected we test for four specifications with the null hypothesises of

(ii) no spatial lag of the dependent variables (vs. a spatial lag),

(iii) no spatial autoregressive error terms (vs . the existence of spatial autoregressive error terms),

(iv) no spatial lag of the dependent variables vs. spatial lag, given spatial autoregressive errors, and finally of

(v) no spatial autoregressive errors vs. the existence of spatial autoregressive errors, given spatial lags of dependent variables.

These (hierarchical) steps are necessary to get indices which of the both models SEMP or SARP should be preferred because both models are not nested and, therefore, could not be directly compared.

The results can be found in Tables A-1A to A-3B for the 140, 138 TTWA and the 406 area data sets. For the 140 area data set the joint test indicate no spatial dependencies for 2 of 4 models including the recession variable. For the model with the full specification and these without the recession variable spatial dependencies cannot be rejected. The further test statistics for the models without recession variables and either no interaction terms for the Kaitz index or no control variables allows us to conclude that the SEMP model should be preferred. For the full specified models (with and without recession variable) the decision between a SARP or SEMP specification is not clear, because in both models the joint tests statistics indicate spatial dependencies but the tests for no spatial lag and no spatial autoregressive term cannot be rejected. The results for the 138 TTWA areas indicate that SEMP would be the preferred model at a significance level of 0.10 with one exception: the SARP model would be preferred for the full specified model with GVA variable[28].

The test results for the 406 areas and the model specification with the commuting matrix indicate for all models that the SEMP should be preferred too. Only for the full specified models with and without recessions variables and with the contiguity matrix the SARP model should be preferred.

[28] However, the p-value is 0.107 and therefore only 0.7 percent points higher than the chosen significance level.

To sum up, the results of the LM tests indicate that spatial dependencies of the dependent variable and/or the error terms cannot be ruled out with a few exceptions. This is especially true for the full specification models. Furthermore, the results lead to the conclusion that the SEMP model should be preferred in the majority of cases. For the other models other spatial specifications should be tested. In such a case LeSage and Pace (2009) recommend to test models with spatial lags of the independent variables that also include spatial lags of independent variables and error terms. However, not least because possible identification problems as it is notably discussed by Gibbons and Overman (2012), it's a task for further research to establish if these spatial specifications are indeed more adequate. Hereby, a contribution to theoretical foundation will be necessary that also will help to derive adequate test statistics and solve the mentioned identification problems. However, we took the result of "a somewhat spatial dependence" into account and applied a further model specification with spatial lags of the Kaitz index. Though, this does not lead to clear inference on the concrete structure of spatial dependence, we are able to compare the results of the estimated coefficients for the employment effects of MW in our model.

[Insert Table A1a here]

[Insert Table A1b here]

[Insert Table A2a here]

[Insert Table A2b here]

[Insert Table A3a here]

[Insert Table A3b here]

D Spatial instruments and specification tests for the SGMM-SLXP model

In order to find a correct model specification with valid instruments that complies with the requirements of the SGMM estimator we conducted several statistical tests.

One requirement for the validity of instruments is that the twice lagged idiosyncratic disturbance term ε_{jt} is not autocorrelated (Arellano and Bond 1991, Roodman 2009a). We use the Arellano-Bond test for autoregression of first and second order in the first differences of the error term. The Null of this test is that there is no autocorrelation. While autocorrelation of first order does not violate the requirements of SGMM and hence the validity of the instruments, the Null of autocorellation of second order should not be rejected. A crucial assumption of this test is that errors are not correlated across cross sectional units. We take this into account in line with Roodman (2009a) since we included time dummy variables as instruments in levels (to handle any cross-sectional dependence in terms of contemporaneous correlation) and furthermore we model structural cross-sectional dependence in terms of spatial lags of the Kaitz index $\rho \sum_{i=1}^{n} w_{ij} P_{it}$. We interpret the coefficient of that term as the nearest parameter for the true direct effect since it can't be fully ruled out that the Kaitz index measured in the observed region is endogenous. Both variables are highly positive correlated, we measured the following correlations[29]:

- commuting matrix: $r_{140\ areas}$=0.7048 ; $r_{138\ areas}$=0.7043; and $r_{406\ areas}$=0.7447;
- contiguity matrix: $r_{140\ areas}$=0.5781 ; $r_{138\ areas}$=0.4544; and $r_{406\ areas}$=0.7052.

An illustration of the correlation can also be found in Figure A-1.

[Insert Figure A-1 about here]

We report all test results in separate tables behind the equivalent results tables of the SGMM estimations, thus in Tables 8, A-5, A-7, A-15, and A-17. The first four rows of those table contain results of the Arellano-Bond test for autoregression of first and second order in the first differences of the error term (test statistics and p-values). In all specification for all geographies Null of no autocorellation of order 1 in first differences of the error term has to be rejected, whereas Null of no autocorrelation of 2nd order could not be rejected.

Hansen's *J* statistic allows to test for the Null of joint validity of all instruments considering that those should be exogenous, thus not correlated with the error term. The test is robust against conditional heteroscedasticity and serial correlation in the error terms (Hansen 1982). Test statistics and p-values can be found in rows 6 and 7 of the mentioned tables. In all specifications for all geographies the Null of joint validity could not be rejected.

Finally, we tested subsets of instruments. There are three groups of instruments. Variables in the first group – GMM instruments – are assumed to be endogenous, hence lags of levels and differences of those instruments are in this group. The employment rate and all or some of the control variables belong to this group. The second group – instrumented variables in levels and first differences – are handled to be strictly exogenous. We assume that the recession variable, the Kaitz index, the spatial lag of the Kaitz index, the incremental difference-in-difference variables, and remaining control variables that are not assigned to the first group are strictly exogenous. Levels and differences of those variables are utilized as instruments. The third group – instrumented variables only in levels – contains only the levels of year effects.

To test for the validity (exogenity) of those subsets we use Difference-in-Hansen or *C*-statistics respectively. They are based on computations of differences of Hansen's *J* statistics for the "unrestricted" model without the subset and the "restricted" model including the subset, thus the increase of *J* statistic after the "unrestricted" model is complemented by the subset (Baum 2006, p. 201

[29] Hereby, all values are significantly unequal zero on a significance level of 1 per cent.

f.). All tests fail to reject the Null of validity of instrument subsets (compare *C*-statistics and p-values in rows 8 to 13 of Tables 8, A-5, A-7, A-15, and A-17).

E Restricted geography sample

We tested the robustness of our results by restricting our samples: we dropped out those regions that are known to be somewhat economically weaker, thus the Western and Northern parts of England, Wales, and Scotland. Therefore we selected 74 from the 140 area sample, 82 from the 138 area sample, and 307 from the 406 area sample. The estimation results can be found in Tables A-8 to A-17.

The results reveal only marginal differences regarding the employment effects of the NMW. The overall effect tends to be zero or negative, whereas the incremental year-by-year effects are more likely to be zero or positive with similar patterns to the results for the full samples. This is also true for the SGMM specification, where the coefficient for the lagged employment rate is – though a bit smaller – significantly positive and thus very robust. The coefficients for the spatial dependency terms get either insignificant or are quite similar to the results for the complete samples. The coefficient for the demand variable reveals the biggest difference: whereas we found a significantly positive influence on employment for each of the full sample data sets, the coefficient tends to be zero or even negative in the case of the restricted samples.

[Insert Table A-8 Here]

[Insert Table A-9 Here]

[Insert Table A-10 Here]

[Insert Table A-11 Here]

[Insert Table A-12 Here]

[Insert Table A-13 Here]

[Insert Table A-14 Here]

[Insert Table A-15 Here]

[Insert Table A-16 Here]

[Insert Table A-17 Here]

F Spatial autocorrelation in the residuals

To assess the residuals for spatial autocorrelation we utilize two tests, the Moran's *I* test and Pessarans *CD* test. The Moran's *I* statistic has the following general form (compare with Greene 2012, Cliff and Ord 1973, Moran 1948):

$$I_t = \frac{n}{\sum_{i=1}^{n}\sum_{j=1}^{n} w_{ij}} \cdot \frac{\sum_{i=1}^{n}\sum_{j=1}^{n} w_{ij}(e_{it}-\overline{e_t})(e_{jt}-\overline{e_t})}{\sum_{i=1}^{n}(e_{it}-\overline{e_t})^2}, i \neq j$$

The terms $e_{it}$and $\overline{e_t}$ denote the residuals for the observation period *t* and region *i* or the mean of the residuals for observations period *t* respectively, whereas the w_{ij} remain to be the *ij*th element of the weights matrix **W**. To consider full panel data the statistic equals $I = \frac{1}{T}\sum_{t=1}^{T} I_t$. Higher values of *I* imply spatial autocorrelation; if *I* is zero there is no spatial autocorrelation in the residuals. An approximation for large samples to the variance of *I* under the null of no spatial autocorrelation is

$$V_I^2 = \frac{1}{T}\frac{n^2\sum_{i=1}^{n}\sum_{j=1}^{n} w_{ij}^2 + 3\left(\sum_{i=1}^{n}\sum_{j=1}^{n} w_{ij}\right)^2 - n\sum_{i=1}^{n}\left(\sum_{j=1}^{n} w_{ij}\right)^2}{(n^2-1)\left(\sum_{i=1}^{n}\sum_{j=1}^{n} w_{ij}\right)^2}$$

Thus, the expected value and variance are a function of the number of independent variables in the system. The term I/V converges to standard normality under the null. That implies that it can be tested if Moran's *I* is significantly different from zero (e.g., the critical value on the 5-per-cent level would be 1,96).

[Insert Table A-18 Here]

[Insert Table A-19 Here]

[Insert Table A-20 Here]

In section II.E we adduced several reasons to prefer a test that does not depends on an assumption about the spatial dependence and rather use Pesaran's *CD* test to test for (remaining) spatial dependence in the residuals. The statistic has the following form

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{T} e_{it}e_{jt}}{(\sum_{t=1}^{T} e_{it}^2)^{1/2}(\sum_{t=1}^{T} e_{jt}^2)^{1/2}}$$

$$CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right)$$

According to Pesaran (2004) *CD* is *N* (0,1) normal-distributed under the null hypothesis of i.i.d. residuals.

[Insert Table A-21 Here]

[Insert Table A-22 Here]

[Insert Table A-23 Here]

Figure A-1. WLA Geography

Source: http://edina.ac.uk/ukborders/

Figure A-2. WAREA Geography

Source: http://edina.ac.uk/ukborders/

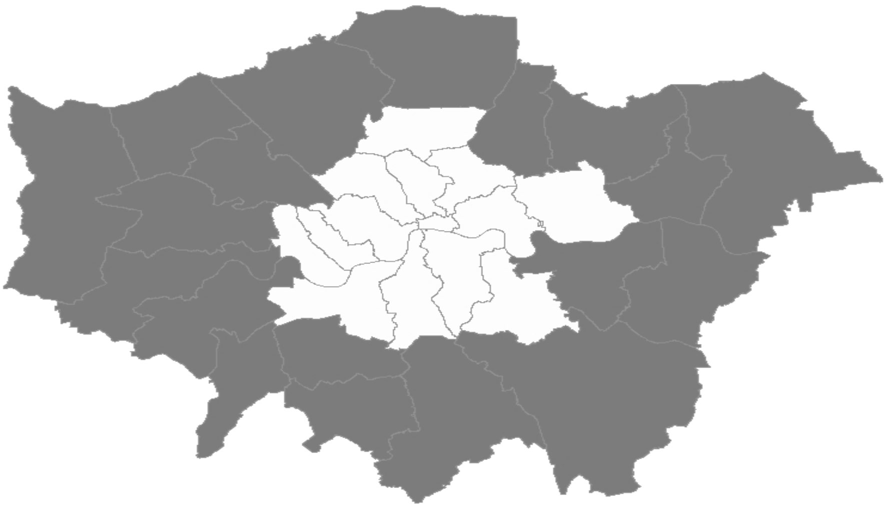


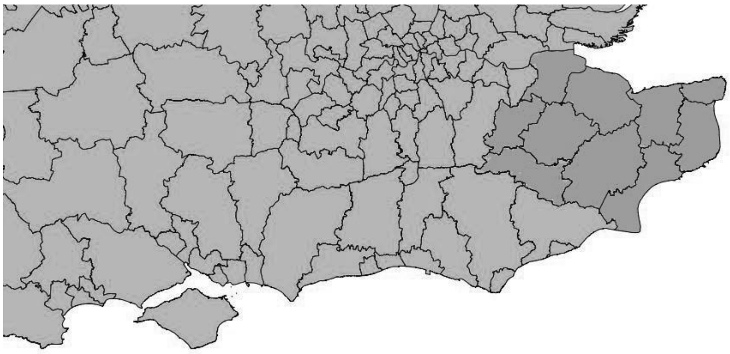
Figure A-3. Differences between WLA and WAREA Geographies, London Area

Source: http://edina.ac.uk/ukborders/

Figure A-4. TTWA Geography

Source: http://edina.ac.uk/ukborders/, geographies are manually created by the authors from WLAs.

a) WLA



b) TTWA

c) WAREA

Figure A-5. Differences between WLA, TTWA and WAREA Geographies, focusing on South-London Coast

Note: The darker regions belong to the county of Kent.

Source: http://edina.ac.uk/ukborders/, TTWA geographies are manually created by the authors from WLAs.

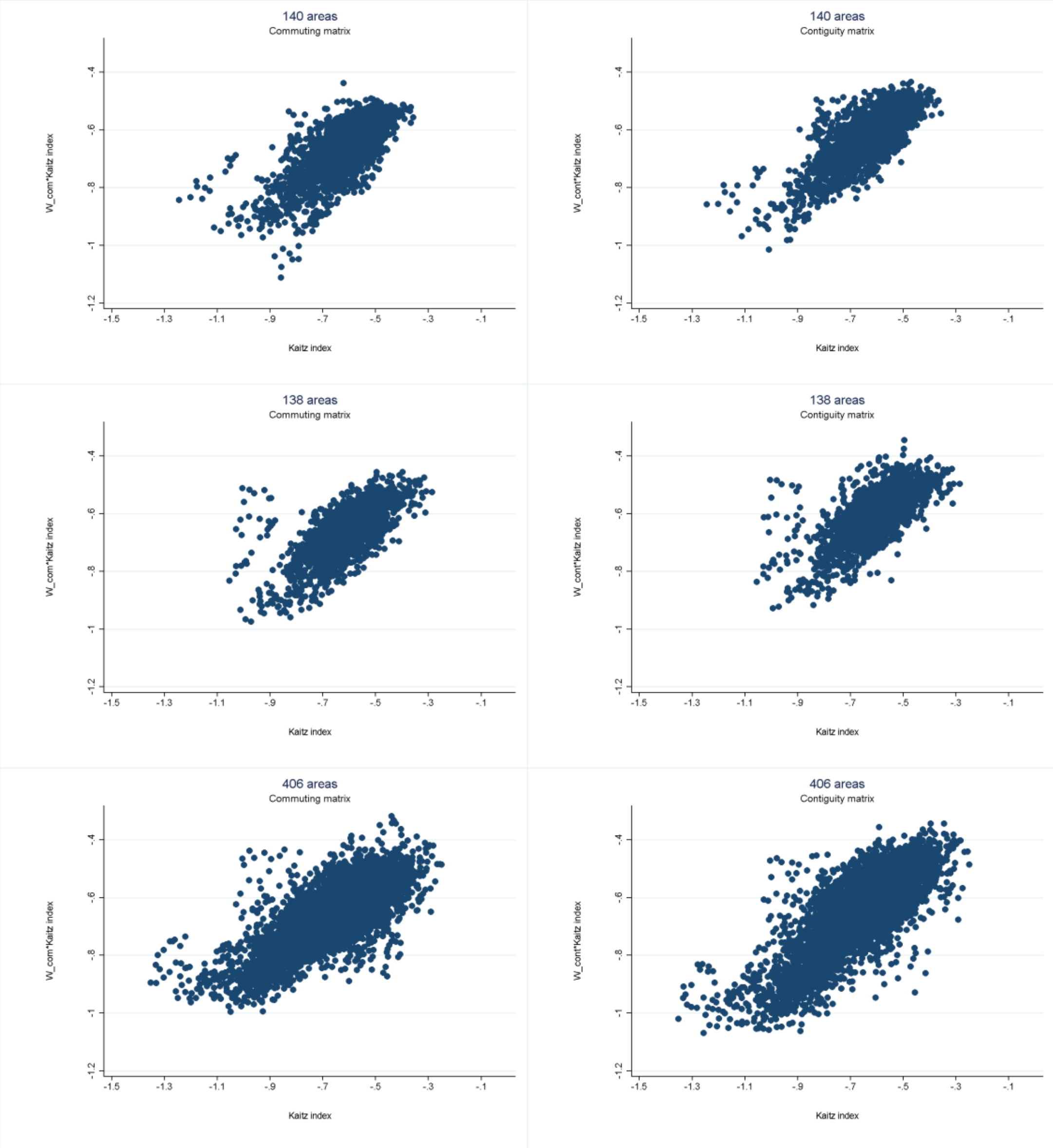


Figure A-6. Scatter plots of (logarithms of) the SLXP terms $\sum_{i=1}^{n} w_{ij}P_{it}$ and the Kaitz index P_{jt} measured in the observed region j (with $i \neq j$).

Note: Axes of Ordinate of the Figures on the left (right) side: SLXP term with commuting (contiguity) matrix.

Table A-1a. LM spatial specification tests, 16 years to retirement age, 140 areas, 1997-2010

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index				Y				
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects				Y				
	Area effects				Y				
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses					Test results				
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	8,984	18,525	16,841	22,246	1,492	5,489	4,142	7,761
	P	0,011	0,000	0,000	0,000	0,474	0,064	0,126	0,021
No spatial lag of the dependent variable	LM statistics	8,938	5,428	4,202	2,499	-	0,399	-	0,093
	P	0,003	0,020	0,040	0,114		0,528		0,760
No spatial autocorrelation of the error terms	LM statistics	8,967	3,767	2,289	0,654	-	0,127	-	0,025
	P	0,003	0,052	0,130	0,419		0,722		0,876
No spatial lag of the dependent variable, given a spatial autocorrelation of the error terms	LM statistics	0,011	0,029	-	-	-	-	-	-
	P	0,915	0,865						
No spatial autocorrelation of the error terms, given a spatial lag of the dependent variable	LM statistics	701,146	220,558	-	-	-	-	-	-
	P	0,000	0,000						
Preferred model (p<0.10)		SEMP	SEMP	SARP	other spec.	no spatial d.	other spec.	no spatial d.	other spec.
	Number of observations				1,960				

Table A-1b. LM spatial specification tests, 16 years to retirement age, 140 areas, 1997-2010, contiguity matrix

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index	Y							
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects	Y							
	Area effects	Y							
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses		Test results							
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	6,531	17,411	14,042	20,366	1,873	5,503	2,440	6,341
	P	0,038	0,000	0,001	0,000	0,392	0,064	0,295	0,042
No spatial lag of the dependent variable	LM statistics	5,268	3,046	1,893	0,987	1,209	0,566	0,426	0,123
	P	0,022	0,081	0,169	0,320	0,272	0,452	0,514	0,726
No spatial autocorrelation of the error terms	LM statistics	4,894	1,456	0,859	0,061	1,026	0,185	0,219	0,003
	P	0,027	0,227	0,354	0,805	0,311	0,667	0,640	0,958
No spatial lag of the dependent variable, given a spatial autocorrelation of the error terms	LM statistics	0,043	-	-	-	-	-	-	-
	P	0,836							
No spatial autocorrelation of the error terms, given a spatial lag of the dependent variable	LM statistics	249,324	-	-	-	-	-	-	-
	P	0,000							
Preferred model (p<0.10)		SEMP	SARP	other spec.	SEMP	no spatial d.	other spec.	no spatial d.	other spec.
	Number of observations	1,960							

Table A-2a. LM spatial specification tests, 16 years to retirement age, 138 TTWA areas, 1997-2010

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index				Y				
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects				Y				
	Area effects				Y				
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses					Test results				
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	14,715	15,708	17,900	17,962	5,658	5,933	7,213	7,281
	P	0,001	0,000	0,000	0,000	0,059	0,051	0,027	0,026
No spatial lag of the dependent variable	LM statistics	12,602	8,850	10,814	7,527	5,284	4,218	4,600	3,605
	P	0,000	0,003	0,001	0,006	0,022	0,040	0,032	0,058
No spatial autocorrelation of the error terms	LM statistics	11,988	7,040	8,952	5,080	4,988	3,507	3,749	2,541
	P	0,001	0,008	0,003	0,024	0,026	0,061	0,053	0,111
No spatial lag of the dependend variable, given a spatial autocorrelation of the error terms	LM statistics	0,017	0,032	0,054	0,223	0,005	0,008	0,006	-
	P	0,895	0,858	0,816	0,637	0,944	0,928	0,936	
No spatial autocorrelation of the error terms, given a spatial lag of the dependend variable	LM statistics	499,611	240,009	314,641	170,207	184,245	107,554	135,457	-
	P	0,000	0,000	0,000	0,000	0,000	0,000	0,000	
Preferred model (p<0.10)		SEMP	SEMP	SEMP	SEMP	SEMP	SEMP	SEMP	SARP
	Number of observations				1,938				

Table A-2b. LM spatial specification tests, 16 years to retirement age, 138 TTWA areas, 1997-2010, contiguity matrix

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index	Y							
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects	Y							
	Area effects	Y							
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses		Test results							
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	8,123	8,691	7,082	8,119	4,278	4,001	3,526	3,365
	p	0,017	0,013	0,029	0,017	0,118	0,135	0,171	0,186
No spatial lag of the dependent variable	LM statistics	7,754	6,388	6,330	5,154	-	-	-	-
	p	0,005	0,011	0,012	0,023				
No spatial autocorrelation of the error terms	LM statistics	7,336	5,004	5,690	3,662	-	-	-	-
	p	0,007	0,025	0,017	0,056				
No spatial lag of the dependent variable, given a spatial autocorrelation of the error terms	LM statistics	0,016	0,077	0,014	0,097	-	-	-	-
	p	0,900	0,781	0,905	0,755				
No spatial autocorrelation of the error terms, given a spatial lag of the dependent variable	LM statistics	181,242	71,509	105,362	55,997	-	-	-	-
	p	0,000	0,000	0,000	0,000				
Preferred model (p<0.10)		SEMP	SEMP	SEMP	SEMP	no spatial d.	no spatial d.	no spatial d.	no spatial d.
	Number of observations	1,938							

Table A-3a. LM spatial specification tests, 16 years to retirement age, 406 areas, 1997-2010

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index	Y							
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects	Y							
	Area effects	Y							
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses		Test results							
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	22,182	23,233	19,977	23,907	21,516	17,864	16,304	18,056
	p	0,000	0,000	0,000	0,000	0,000	0,000	0,000	0,000
No spatial lag of the dependent variable	LM statistics	20,790	19,399	14,482	13,309	12,667	12,275	9,974	9,517
	p	0,000	0,000	0,000	0,000	0,000	0,000	0,002	0,002
No spatial autocorrelation of the error terms	LM statistics	20,547	17,547	13,118	10,392	11,552	10,410	8,797	7,350
	p	0,000	0,000	0,000	0,001	0,001	0,001	0,003	0,007
No spatial lag of the dependent variable, given a spatial autocorrelation of the error terms	LM statistics	0,090	0,038	0,023	0,080	0,017	0,025	0,052	0,212
	p	0,764	0,845	0,879	0,777	0,897	0,875	0,819	0,645
No spatial autocorrelation of the error terms, given a spatial lag of the dependent variable	LM statistics	1794,738	598,331	781,769	347,867	1144,551	369,221	580,930	274,090
	p	0,000	0,000	0,000	0,000	0,000	0,000	0,000	0,000
Preferred model (p<0.10)		SEMP	SEMP	SEMP	SEMP	SEMP	SEMP	SEMP	SEMP
	Number of observations	5,684							

Table A-3b. LM spatial specification tests, 16 years to retirement age, 406 areas, 1997-2010, contiguity matrix

	Specification of the model	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Specification of the model	Kaitz index					Y			
	Kaitz*Year effects	N	N	Y	Y	N	N	Y	Y
	Year effects					Y			
	Area effects					Y			
	Controls incl. share publ.	N	Y	N	Y	N	Y	N	Y
	GVA	N	N	N	N	Y	Y	Y	Y
Null hypotheses						Test results			
No spatial autocorrelation of the error terms and no spatial lag of the dependent variables	LM statistics	17,040	12,990	19,125	18,012	15,858	8,585	14,061	11,472
	P	0,000	0,002	0,000	0,000	0,000	0,014	0,001	0,003
No spatial lag of the dependent variable	LM statistics	9,281	7,889	5,484	4,422	5,384	4,654	3,611	2,962
	P	0,002	0,005	0,019	0,035	0,020	0,031	0,057	0,085
No spatial autocorrelation of the error terms	LM statistics	8,561	6,327	4,120	2,458	4,573	3,600	2,701	1,728
	P	0,003	0,012	0,042	0,117	0,032	0,058	0,100	0,189
No spatial lag of the dependent variable, given a spatial autocorrelation of the error terms	LM statistics	0,040	0,054	0,137	0,327	0,074	0,041	-	-
	P	0,842	0,816	0,712	0,568	0,786	0,840		
No spatial autocorrelation of the error terms, given a spatial lag of the dependent variable	LM statistics	1247,100	207,153	313,893	98,987	581,534	113,083	-	-
	P	0,000	0,000	0,000	0,000	0,000	0,000		
Preferred model (p<0.10)		SEMP	SEMP	SEMP	SARP	SEMP	SEMP	SARP	SARP
	Number of observations					5,684			

Table A-4. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 1997-2010, all regressions contain control variables, area and year effects..

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions				138 regions				406 regions			
	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Contiguity	SGMM SLXP Contiguity
Kaitz Index	-0.006	0.023	-0.050	-0.024	-0.104	-0.069	-0.120	0.055	-0.002	0.006	-0.028	-0.017
	(0.066)	(0.066)	(0.068)	(0.075)	(0.069)	(0.072)	(0.087)	(0.060)	(0.038)	(0.038)	(0.037)	(0.039)
w*Kaitz Index	-0.173***	-0.159**	-0.111**	-0.101*	-0.243***	-0.245***	-0.106	-0.159***	-0.106**	-0.105**	-0.021	-0.019
	(0.062)	(0.061)	(0.052)	(0.053)	(0.065)	(0.068)	(0.083)	(0.053)	(0.044)	(0.043)	(0.038)	(0.037)
E_{t-1}	0.259***	0.251***	0.252***	0.247***	0.235***	0.224***	0.240***	0.253***	0.228***	0.227***	0.230***	0.229***
	(0.032)	(0.032)	(0.035)	(0.035)	(0.042)	(0.042)	(0.040)	(0.034)	(0.024)	(0.024)	(0.024)	(0.024)
GVA		0.323		0.506**		0.931***		0.497		0.148		0.210
		(0.258)		(0.250)		(0.317)		(0.301)		(0.192)		(0.186)
Share Public	-0.141***	-0.149***	-0.180***	-0.184***	-0.081	-0.082	-0.157	-0.151***	-0.181***	-0.181***	-0.213***	-0.215***
	(0.041)	(0.041)	(0.038)	(0.038)	(0.050)	(0.050)	(0.107)	(0.036)	(0.030)	(0.030)	(0.028)	(0.028)
Kaitz*1999	0.006	0.006	0.016	0.021	0.060	0.057	0.042	0.056	-0.034	-0.036	-0.033	-0.037
	(0.051)	(0.052)	(0.057)	(0.060)	(0.059)	(0.059)	(0.049)	(0.059)	(0.026)	(0.026)	(0.026)	(0.026)
Kaitz*2000	0.045	0.039	0.050	0.040	0.026	0.007	0.018	0.026	0.012	0.006	0.017	0.009
	(0.048)	(0.049)	(0.053)	(0.055)	(0.052)	(0.053)	(0.054)	(0.049)	(0.024)	(0.026)	(0.025)	(0.026)
Kaitz*2001	-0.049	-0.064	-0.049	-0.068	-0.052	-0.076	-0.059	-0.085	-0.028	-0.037	-0.026	-0.038
	(0.040)	(0.042)	(0.048)	(0.050)	(0.047)	(0.046)	(0.044)	(0.055)	(0.025)	(0.028)	(0.025)	(0.028)
Kaitz*2002	-0.018	-0.038	-0.013	-0.039	0.022	-0.015	0.007	-0.022	0.005	-0.004	0.012	-0.002
	(0.040)	(0.042)	(0.047)	(0.049)	(0.045)	(0.047)	(0.042)	(0.043)	(0.026)	(0.029)	(0.026)	(0.029)
Kaitz*2003	0.127**	0.105*	0.131*	0.101	0.086*	0.049	0.053	-0.024	0.044	0.036	0.046	0.035
	(0.055)	(0.056)	(0.068)	(0.070)	(0.052)	(0.051)	(0.051)	(0.057)	(0.031)	(0.032)	(0.030)	(0.031)
Kaitz*2004	0.081	0.063	0.073	0.037	0.042	0.006	0.023	0.046	0.024	0.016	0.025	0.013
	(0.056)	(0.061)	(0.059)	(0.064)	(0.074)	(0.074)	(0.073)	(0.072)	(0.029)	(0.030)	(0.029)	(0.030)
Kaitz*2005	0.090**	0.076	0.068	0.043	-0.002	-0.039	-0.018	-0.021	0.020	0.011	0.022	0.010
	(0.045)	(0.048)	(0.052)	(0.056)	(0.062)	(0.064)	(0.068)	(0.056)	(0.034)	(0.036)	(0.034)	(0.036)
Kaitz*2006	0.102**	0.080	0.119**	0.078	-0.007	-0.045	-0.032	-0.053	0.019	0.010	0.022	0.010
	(0.045)	(0.051)	(0.056)	(0.060)	(0.060)	(0.061)	(0.060)	(0.061)	(0.039)	(0.041)	(0.039)	(0.041)
Kaitz*2007	0.031	0.013	0.034	0.002	-0.010	-0.043	-0.038	0.035	-0.001	-0.010	0.006	-0.006
	(0.057)	(0.063)	(0.067)	(0.070)	(0.073)	(0.073)	(0.065)	(0.080)	(0.035)	(0.037)	(0.035)	(0.037)
Kaitz*2008	0.061	0.044	0.094*	0.059	0.103	0.071	0.047	0.003	-0.012	-0.021	-0.006	-0.017
	(0.050)	(0.056)	(0.053)	(0.058)	(0.064)	(0.064)	(0.056)	(0.057)	(0.036)	(0.038)	(0.036)	(0.038)
Kaitz*2009	-0.006	-0.019	0.031	-0.006	-0.008	-0.040	0.000	-0.135*	-0.039	-0.049	-0.025	-0.038
	(0.050)	(0.057)	(0.055)	(0.061)	(0.067)	(0.069)	(0.063)	(0.070)	(0.039)	(0.043)	(0.039)	(0.042)
Kaitz*2010	0.033	0.012	0.066	0.018	0.050	0.001	0.065	-0.025	0.064	0.054	0.074*	0.059
	(0.061)	(0.067)	(0.072)	(0.075)	(0.069)	(0.072)	(0.065)	(0.077)	(0.041)	(0.042)	(0.040)	(0.042)
Observations	1,820	1,820	1,781	1,781	1,794	1,794	1,742	1,742	5,278	5,278	5,213	5,213
No. of instruments	70	71	70	71	70	71	70	71	70	71	70	71

*** p<0.01, ** p<0.05, * p<0.1. Test statistics are provided in the next Table.
In SLXP Models with contiguity matrix the islands are excluded.

Table A-5. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, excluding direct effect of Minimum Wage, test statistics of estimates in table A-4.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions				138 regions				406 regions			
	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA
Number of instruments	70	71	70	71	70	71	70	71	70	71	70	71
Arellano-Bond test for AR in first differences												
AR(1)	-6.3840	-6.3719	-6.1274	-6.1491	-6.3902	-6.3973	-6.7373	-6.9406	-13.9528	-13.9400	-13.9678	-13.9558
Prob > z	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
AR(2)	1.0516	1.0281	0.7214	0.7103	-0.5654	-0.5878	-1.3688	-1.4117	-0.3149	-0.3119	-0.6288	-0.6369
Prob > z	0.2930	0.3039	0.4706	0.4775	0.5718	0.5566	0.1711	0.1580	0.7529	0.7551	0.5295	0.5242
Hansen test of overidentified restrictions												
J	29.1365	30.5191	37.6739	37.9516	42.1832	43.9875	39.6803	42.9119	40.2840	39.8107	39.4990	38.7964
Prob> chi2	0.7843	0.7265	0.3926	0.3805	0.2211	0.1693	0.3093	0.1990	0.2864	0.3043	0.3164	0.3447
Difference-in-Hansen tests of exogeneity of instrument subsets												
- GMM instruments for levels -												
C	27.7384	28.0837	34.5093	35.0653	41.6844	42.7971	38.2968	41.1118	36.1659	35.7353	33.6713	33.0885
Prob > chi2	0.7265	0.7105	0.3955	0.3704	0.1428	0.1182	0.2416	0.1569	0.3230	0.3411	0.4348	0.4629
- Instrumented variables in levels and first differences -												
C	16.5076	16.2332	20.2653	19.6337	23.3270	21.3813	23.2171	18.3432	25.1586	21.3571	24.5215	22.6525
Prob > chi2	0.5572	0.5074	0.3181	0.2934	0.1783	0.2097	0.1824	0.3675	0.1206	0.2107	0.1387	0.1609
- Instrumented variables only in levels -												
C	19.9169	20.3446	25.1119	25.0303	24.5003	26.8617	32.2114	33.1708	25.6343	25.2604	24.9355	24.3277
Prob > chi2	0.7015	0.6770	0.3997	0.4041	0.4333	0.3109	0.1218	0.1005	0.3720	0.3917	0.4093	0.4430

Table A-6. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, excluding direct effect of Minimum Wage, 16 years to retirement age, 1997-2010, all regressions contain control variables, area and year effects. Contiguity Matrix

	(1)	(2)	(3)	(4)	(5)	(6)
	140 regions		138 regions		406 regions	
	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Contiguity
Kaitz Index	-	-	-	-	-	-
	-	-	-	-	-	-
w*Kaitz Index	-0.095	-0.095	-0.150**	-0.100*	-0.029	-0.025
	(0.069)	(0.065)	(0.067)	(0.052)	(0.035)	(0.041)
E_{t-1}	0.244***	0.241***	0.239***	0.224***	0.229***	0.221***
	(0.034)	(0.034)	(0.041)	(0.032)	(0.024)	(0.024)
GVA		0.255		0.496		0.242
		(0.316)		(0.331)		(0.184)
Share Public	-0.272**	-0.261**	-0.113	-0.251***	-0.211***	-0.210***
	(0.108)	(0.104)	(0.093)	(0.083)	(0.028)	(0.057)
Kaitz*1999	0.003	0.014	-0.009	-0.001	-0.048*	-0.052*
	(0.045)	(0.044)	(0.060)	(0.053)	(0.028)	(0.029)
Kaitz*2000	0.047	0.056	-0.053	-0.006	-0.003	0.000
	(0.057)	(0.055)	(0.064)	(0.055)	(0.030)	(0.035)
Kaitz*2001	-0.053	-0.045	-0.132*	-0.113	-0.048	-0.048
	(0.056)	(0.055)	(0.072)	(0.071)	(0.030)	(0.037)
Kaitz*2002	-0.011	-0.007	-0.069	-0.045	-0.010	-0.017
	(0.063)	(0.062)	(0.071)	(0.070)	(0.034)	(0.040)
Kaitz*2003	0.088	0.090	-0.033	-0.028	0.023	0.020
	(0.057)	(0.057)	(0.061)	(0.067)	(0.028)	(0.036)
Kaitz*2004	0.062	0.064	-0.061	0.005	0.002	0.006
	(0.073)	(0.072)	(0.077)	(0.083)	(0.034)	(0.042)
Kaitz*2005	0.071	0.074	-0.098	-0.018	0.000	0.009
	(0.072)	(0.071)	(0.087)	(0.082)	(0.039)	(0.044)
Kaitz*2006	0.078	0.076	-0.110	-0.067	0.002	-0.008
	(0.069)	(0.068)	(0.083)	(0.086)	(0.048)	(0.052)
Kaitz*2007	0.011	0.012	-0.114	-0.004	-0.019	-0.021
	(0.080)	(0.080)	(0.084)	(0.081)	(0.039)	(0.042)
Kaitz*2008	0.048	0.048	-0.036	-0.008	-0.029	-0.015
	(0.071)	(0.070)	(0.078)	(0.078)	(0.038)	(0.041)
Kaitz*2009	0.008	0.008	-0.086	-0.136*	-0.047	-0.039
	(0.079)	(0.079)	(0.077)	(0.077)	(0.046)	(0.050)
Kaitz*2010	0.074	0.073	-0.016	-0.058	0.051	0.057
	(0.070)	(0.070)	(0.080)	(0.077)	(0.044)	(0.050)
Observations	1,781	1,781	1,742	1,742	5,213	5,213
Number of instruments	69	70	69	82	69	70

*** p<0.01, ** p<0.05, * p<0.1. In SLXP Models with contiguity matrix the islands are excluded.

Table A-7. SGMM-SLXP Estimates of Minimum Wage Effects on Employment, test statistics of estimates in table A-6

	(1)	(2)	(3)	(4)	(5)	(6)
	140 regions		138 regions		406 regions	
	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA
Number of instruments	69	70	69	82	69	70
	Arellano-Bond test for AR in first differences					
AR(1)	-6.1772	-6.1636	-6.6484	-6.7160	-13.9372	-13.8655
Prob > z	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
AR(2)	0.7527	0.7550	-1.4377	-1.4724	-0.6322	-0.7134
Prob > z	0.4516	0.4502	0.1505	0.1409	0.5272	0.4756
	Hansen test of overidentified restrictions					
J	36.7006	36.9406	40.5304	55.4910	39.6551	38.0651
Prob> chi2	0.4362	0.4253	0.2773	0.2132	0.3103	0.3756
	Difference-in-Sargan tests of exogeneity of instrument subsets					
	- GMM instruments -					
C	32.4007	32.9656	39.4155	49.3487	33.8544	34.2945
Prob > chi2	0.4968	0.4689	0.2048	0.2679	0.4261	0.4055
	- Instrumented variables in levels and first differences -					
C	18.1875	18.3662	24.3882	31.7669	24.6164	17.6435
Prob > chi2	0.5099	0.4318	0.1817	0.4281	0.1736	0.4794
	- Instrumented variables only in levels -					
C	28.8001	29.0212	32.5703	41.5760	25.0771	25.1066
Prob > chi2	0.2277	0.2193	0.1135	0.2408	0.4016	0.4000

Table A-8. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 140 areas level (74 selected regions), 1997-2010, all regressions contain control variables, area and year effects.

	(1) FE	(2) SEMP Commuting Matrix	(3) SEMP Contiguity Matrix	(4) SLXP Commuting Matrix	(5) SLXP Contiguity Matrix	(6) FE	(7) SEMP Commuting Matrix	(8) SEMP Contiguity Matrix	(9) SLXP Commuting Matrix	(10) SLXP Contiguity Matrix
Kaitz Index	-0.090	-0.090**	-0.090**	-0.090	-0.090	-0.097*	-0.096**	-0.096**	-0.096*	-0.096*
	(0.056)	(0.038)	(0.038)	(0.055)	(0.054)	(0.055)	(0.038)	(0.038)	(0.055)	(0.054)
w*Kaitz Index				0.281	0.009				0.194	-0.018
				(0.207)	(0.213)				(0.214)	(0.205)
GVA						0.433*	0.428**	0.434**	0.379	0.437*
						(0.226)	(0.209)	(0.211)	(0.235)	(0.231)
Share Public	-0.050	-0.049	-0.049	-0.049	-0.050	-0.047	-0.047	-0.047	-0.047	-0.047
	(0.043)	(0.041)	(0.041)	(0.042)	(0.043)	(0.043)	(0.041)	(0.041)	(0.042)	(0.043)
age	-2.111	-2.059	-1.966	-2.012	-2.121	-1.783	-1.75	-1.651	-1.756	-1.760
	(1.986)	(2.690)	(2.691)	(1.994)	(1.999)	(2.038)	(2.690)	(2.690)	(2.041)	(2.057)
age2	0.051	0.049	0.047	0.048	0.051	0.042	0.041	0.039	0.042	0.042
	(0.049)	(0.068)	(0.068)	(0.050)	(0.050)	(0.051)	(0.068)	(0.068)	(0.051)	(0.051)
age3	-0.000	0.000	0.000	-0.000	-0.000	-0.000	0.000	0.000	-0.000	-0.000
	(0.000)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.000)	(0.000)
nvq4plusIMP	0.159***	0.158***	0.155***	0.157***	0.159***	0.162***	0.161***	0.159***	0.160***	0.162***
	(0.046)	(0.041)	(0.041)	(0.047)	(0.046)	(0.046)	(0.041)	(0.041)	(0.046)	(0.046)
total_female	-0.021	-0.024	-0.024	-0.025	-0.021	-0.024	-0.026	-0.026	-0.026	-0.023
	(0.081)	(0.067)	(0.067)	(0.079)	(0.080)	(0.081)	(0.067)	(0.067)	(0.079)	(0.080)
lambda		0.053	0.054				0.036	0.049		
		(0.051)	(0.041)				(0.051)	(0.041)		
Observations	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036
R-squared	0.182			0.184	0.182	0.187			0.187	0.186
ll		2,042.279	2,042.588				2,044.514	2,044.839		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10)

*** p<0.01, ** p<0.05, * p<0.1

Table A-9. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 138 areas level (82 selected regions), 1997-2010, all regressions contain control variables, area and year effects.

	(1) FE	(2) SEMP Commuting Matrix	(3) SEMP Contiguity Matrix	(4) SLXP Commuting Matrix	(5) SLXP Contiguity Matrix	(6) FE	(7) SEMP Commuting Matrix	(8) SEMP Contiguity Matrix	(9) SLXP Commuting Matrix	(10) SLXP Contiguity Matrix
Kaitz Index	-0.021	-0.023	-0.021	-0.028	-0.027	-0.021	-0.022	-0.021	-0.028	-0.026
	(0.045)	(0.045)	(0.045)	(0.039)	(0.042)	(0.045)	(0.045)	(0.045)	(0.040)	(0.042)
w^{com}*Kaitz Index				0.092	0.073				0.097	0.077
				(0.158)	(0.108)				(0.161)	(0.110)
GVA						-0.050	-0.065	-0.033	-0.080	-0.082
						(0.302)	(0.258)	(0.265)	(0.310)	(0.306)
Share Public	-0.088	-0.089**	-0.088**	-0.088*	-0.087	-0.087	-0.089**	-0.088**	-0.088*	-0.087
	(0.053)	(0.043)	(0.043)	(0.053)	(0.053)	(0.053)	(0.043)	(0.043)	(0.052)	(0.053)
age	6.103*	5.915**	6.075**	6.131*	6.117*	6.083*	5.894**	6.069**	6.101*	6.086*
	(3.114)	(2.589)	(2.545)	(3.103)	(3.100)	(3.099)	(2.590)	(2.546)	(3.084)	(3.078)
age2	-0.150*	-0.145**	-0.149**	-0.151*	-0.150*	-0.149*	-0.145**	-0.149**	-0.150*	-0.149*
	(0.077)	(0.064)	(0.063)	(0.077)	(0.077)	(0.077)	(0.064)	(0.063)	(0.077)	(0.077)
age3	0.001*	0.001**	0.001**	0.001*	0.001*	0.001*	0.001**	0.001**	0.001*	0.001*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
nvq4plusIMP	0.192***	0.196***	0.200***	0.193***	0.194***	0.192***	0.196***	0.200***	0.194***	0.195***
	(0.049)	(0.036)	(0.036)	(0.049)	(0.049)	(0.049)	(0.036)	(0.036)	(0.049)	(0.049)
total_female	-0.038	-0.035	-0.034	-0.034	-0.037	-0.039	-0.035	-0.034	-0.034	-0.037
	(0.063)	(0.065)	(0.065)	(0.063)	(0.063)	(0.063)	(0.065)	(0.065)	(0.063)	(0.063)
lambda		0.069	0.089**				0.069	0.091**		
		(0.049)	(0.041)				(0.049)	(0.041)		
Observations	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148
R-squared	0.180			0.180	0.180	0.180			0.180	0.180
ll		2050,113	2,051.735				2,050.145	2,051.744		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10)

*** p<0.01, ** p<0.05, * p<0.1

Table A-10. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 406 areas level (307 selected regions), 1997-2010, all regressions contain control variables, area and year effects.

	(1) FE	(2) SEMP Commuting Matrix	(3) SEMP Contiguity Matrix	(4) SLXP Commuting Matrix	(5) SLXP Contiguity Matrix	(6) FE	(7) SEMP Commuting Matrix	(8) SEMP Contiguity Matrix	(9) SLXP Commuting Matrix	(10) SLXP Contiguity Matrix
Kaitz Index	-0.014	-0,014	-0,014	-0.015	-0.015	-0.014	-0,014	-0,014	-0.015	-0.015
	(0.026)	(0.022)	(0.022)	(0.026)	(0.026)	(0.026)	(0.022)	(0.022)	(0.026)	(0.026)
w^{com}*Kaitz Index				0.060	0.039				0.056	0.037
				(0.115)	(0.088)				(0.117)	(0.089)
GVA						0.054	0,026	0,043	0.043	0.048
						(0.171)	(0.154)	(0.151)	(0.174)	(0.172)
Share Public	0.005	0,004	0,004	0.005	0.004	0.004	0,004	0,004	0.004	0.004
	(0.028)	(0.025)	(0.025)	(0.028)	(0.028)	(0.028)	(0.025)	(0.025)	(0.028)	(0.028)
age	2.658	2,701***	2,668***	2.684	2.671	2.668	2,708***	2,679***	2.691	2.679
	(1.675)	(0.994)	(0.995)	(1.677)	(1.682)	(1.675)	(0.995)	(0.995)	(1.677)	(1.682)
age2	-0.067	-0,068***	-0,067***	-0.067	-0.067	-0.067	-0,068***	-0,067***	-0.068	-0.067
	(0.042)	(0.025)	(0.025)	(0.042)	(0.042)	(0.042)	(0.025)	(0.025)	(0.042)	(0.042)
age3	0.001	0,001***	0,001***	0.001	0.001	0.001	0,001***	0,001***	0.001	0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
nvq4plusIMP	0.172***	0,172***	0,171***	0.172***	0.172***	0.172***	0,172***	0,171***	0.172***	0.172***
	(0.026)	(0.019)	(0.019)	(0.026)	(0.026)	(0.026)	(0.019)	(0.019)	(0.026)	(0.026)
total_female	0.024	0,023	0,024	0.025	0.025	0.025	0,023	0,024	0.025	0.025
	(0.038)	(0.034)	(0.034)	(0.038)	(0.038)	(0.038)	(0.034)	(0.034)	(0.038)	(0.038)
lambda		0,068**	0,032				0,070**	0,035		
		(0.029)	(0.023)				(0.029)	(0.023)		
Observations	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298
R-squared	0.083			0.083	0.083	0.083			0.083	0.083
ll		6,283.053	6,281.658				6283,07	6,281.71		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10)

*** p<0.01, ** p<0.05, * p<0.1

Table A-11. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 140 areas level (74 selected regions), 1997-2010, all regressions contain control variables, area and year effects.

	(1) FE	(2) SEMP Commuting	(3) SEMP Contiguity	(4) SLXP Commuting	(5) SLXP Contiguity	(6) FE	(7) SEMP Commuting	(8) SEMP Contiguity	(9) SLXP Commuting	(10) SLXP Contiguity
Kaitz Index	-0.171***	-0.171***	-0.171***	-0.170***	-0.171***	-0.171***	-0.171***	-0.171***	-0.170***	-0.170***
	(0.061)	(0.046)	(0.046)	(0.061)	(0.060)	(0.061)	(0.046)	(0.046)	(0.061)	(0.060)
w^{com}*Kaitz Index				0.239	-0.004				0.200	-0.019
				(0.219)	(0.212)				(0.223)	(0.205)
GVA						0.255	0.255	0.256	0.206	0.259
						(0.255)	(0.213)	(0.216)	(0.260)	(0.251)
Share Public	-0.046	-0.046	-0.046	-0.045	-0.046	-0.045	-0.045	-0.045	-0.045	-0.045
	(0.043)	(0.041)	(0.041)	(0.043)	(0.043)	(0.044)	(0.041)	(0.041)	(0.043)	(0.043)
Kaitz*1999	0.003	0.003	0.003	0.001	0.003	0.006	0.006	0.006	0.004	0.007
	(0.035)	(0.042)	(0.042)	(0.035)	(0.035)	(0.035)	(0.042)	(0.042)	(0.035)	(0.035)
Kaitz*2000	0.084**	0.084**	0.085**	0.086**	0.085**	0.086**	0.085**	0.086**	0.087**	0.086**
	(0.034)	(0.043)	(0.043)	(0.033)	(0.034)	(0.034)	(0.042)	(0.043)	(0.034)	(0.034)
Kaitz*2001	0.071	0.071*	0.071*	0.072*	0.071	0.068	0.067*	0.067	0.069	0.068
	(0.044)	(0.041)	(0.041)	(0.043)	(0.044)	(0.046)	(0.041)	(0.041)	(0.045)	(0.046)
Kaitz*2002	0.094**	0.094**	0.094**	0.092**	0.094**	0.087**	0.087**	0.087**	0.087**	0.087**
	(0.036)	(0.041)	(0.041)	(0.036)	(0.037)	(0.039)	(0.041)	(0.042)	(0.039)	(0.040)
Kaitz*2003	0.159**	0.159***	0.158***	0.159**	0.159**	0.153**	0.153***	0.152***	0.154**	0.153**
	(0.065)	(0.042)	(0.043)	(0.064)	(0.066)	(0.067)	(0.042)	(0.043)	(0.066)	(0.067)
Kaitz*2004	0.114**	0.114***	0.114***	0.112**	0.114**	0.108**	0.108**	0.108**	0.107**	0.108**
	(0.050)	(0.042)	(0.043)	(0.050)	(0.051)	(0.052)	(0.042)	(0.043)	(0.051)	(0.052)
Kaitz*2005	0.112***	0.111**	0.112**	0.106***	0.112***	0.106***	0.105**	0.106**	0.102***	0.106***
	(0.035)	(0.044)	(0.044)	(0.037)	(0.036)	(0.037)	(0.044)	(0.044)	(0.038)	(0.038)
Kaitz*2006	0.152***	0.152***	0.152***	0.146***	0.152***	0.146***	0.146***	0.146***	0.142***	0.146***
	(0.045)	(0.043)	(0.043)	(0.044)	(0.045)	(0.046)	(0.043)	(0.044)	(0.046)	(0.046)
Kaitz*2007	0.065	0.065	0.064	0.058	0.065	0.059	0.059	0.059	0.055	0.059
	(0.052)	(0.043)	(0.043)	(0.051)	(0.052)	(0.052)	(0.043)	(0.043)	(0.052)	(0.052)
Kaitz*2008	0.097**	0.097**	0.096**	0.090*	0.097**	0.089*	0.089**	0.088**	0.084*	0.089*
	(0.046)	(0.042)	(0.042)	(0.046)	(0.047)	(0.046)	(0.042)	(0.043)	(0.046)	(0.047)
Kaitz*2009	0.037	0.037	0.037	0.031	0.037	0.029	0.029	0.029	0.026	0.029
	(0.050)	(0.043)	(0.043)	(0.051)	(0.051)	(0.052)	(0.044)	(0.044)	(0.052)	(0.052)
Kaitz*2010	0.100*	0.100**	0.100**	0.094*	0.100*	0.088	0.088**	0.088*	0.085	0.088
	(0.050)	(0.044)	(0.044)	(0.051)	(0.051)	(0.054)	(0.045)	(0.045)	(0.054)	(0.054)
lambda		-0.011	0.007				-0.015	0.007		
		(0.052)	(0.041)				(0.052)	(0.041)		
Observations	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036	1,036
R-squared	0.206			0.207	0.206	0.207			0.208	0.207
ll		2,056.94	2,056.941				2,057.709	2,057.697		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10)*** p<0.01, ** p<0.05, * p<0.1

Table A-12. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 138 areas level (82 selected regions), 1997-2010, all regressions contain control variables, area and year effects.

	(1) FE	(2) SEMP Commuting	(3) SEMP Contiguity	(4) SLXP Commuting	(5) SLXP Contiguity	(6) FE	(7) SEMP Commuting	(8) SEMP Contiguity	(9) SLXP Commuting	(10) SLXP Contiguity
Kaitz Index	-0.057	-0.057	-0.056	-0.071	-0.069	-0.057	-0.057	-0.055	-0.072	-0.069
	(0.049)	(0.053)	(0.054)	(0.045)	(0.047)	(0.049)	(0.053)	(0.054)	(0.046)	(0.048)
w*Kaitz Index				0.143	0.116				0.169	0.135
				(0.153)	(0.107)				(0.155)	(0.110)
GVA						-0.242	-0.246	-0.222	-0.311	-0.326
						(0.334)	(0.261)	(0.268)	(0.336)	(0.331)
Share Public	-0.088	-0.089**	-0.088**	-0.090*	-0.087	-0.087	-0.088**	-0.087**	-0.088*	-0.085
	(0.053)	(0.043)	(0.043)	(0.053)	(0.054)	(0.053)	(0.043)	(0.043)	(0.053)	(0.053)
Kaitz*1999	-0.034	-0.035	-0.037	-0.030	-0.031	-0.036	-0.037	-0.039	-0.032	-0.034
	(0.055)	(0.053)	(0.054)	(0.057)	(0.055)	(0.055)	(0.053)	(0.054)	(0.057)	(0.055)
Kaitz*2000	0.041	0.04	0.037	0.043	0.042	0.039	0.038	0.036	0.041	0.039
	(0.039)	(0.056)	(0.057)	(0.041)	(0.040)	(0.040)	(0.056)	(0.057)	(0.041)	(0.041)
Kaitz*2001	-0.012	-0.013	-0.012	-0.012	-0.013	-0.009	-0.01	-0.009	-0.008	-0.009
	(0.048)	(0.054)	(0.055)	(0.049)	(0.049)	(0.049)	(0.054)	(0.055)	(0.050)	(0.051)
Kaitz*2002	0.029	0.027	0.025	0.029	0.029	0.035	0.033	0.031	0.037	0.037
	(0.037)	(0.053)	(0.055)	(0.037)	(0.037)	(0.039)	(0.054)	(0.055)	(0.038)	(0.038)
Kaitz*2003	0.101**	0.099*	0.095*	0.106**	0.105**	0.106**	0.104*	0.1*	0.113**	0.113**
	(0.047)	(0.054)	(0.055)	(0.048)	(0.047)	(0.048)	(0.054)	(0.055)	(0.049)	(0.048)
Kaitz*2004	0.129***	0.128**	0.128**	0.134***	0.134***	0.134***	0.133**	0.133**	0.142***	0.142***
	(0.044)	(0.055)	(0.056)	(0.044)	(0.045)	(0.046)	(0.055)	(0.056)	(0.045)	(0.046)
Kaitz*2005	0.069	0.066	0.063	0.073	0.074	0.073	0.07	0.066	0.079	0.080
	(0.047)	(0.055)	(0.057)	(0.047)	(0.048)	(0.049)	(0.056)	(0.057)	(0.049)	(0.050)
Kaitz*2006	0.024	0.02	0.014	0.032	0.031	0.027	0.023	0.018	0.038	0.037
	(0.069)	(0.058)	(0.059)	(0.070)	(0.071)	(0.069)	(0.058)	(0.059)	(0.071)	(0.072)
Kaitz*2007	0.053	0.049	0.041	0.062	0.062	0.057	0.052	0.044	0.068*	0.069*
	(0.038)	(0.058)	(0.059)	(0.040)	(0.040)	(0.039)	(0.058)	(0.059)	(0.040)	(0.040)
Kaitz*2008	0.086*	0.086	0.086	0.090*	0.094*	0.092*	0.092	0.091	0.099*	0.103**
	(0.047)	(0.056)	(0.057)	(0.048)	(0.048)	(0.049)	(0.057)	(0.058)	(0.050)	(0.051)
Kaitz*2009	0.020	0.022	0.028	0.029	0.032	0.027	0.028	0.034	0.039	0.043
	(0.059)	(0.059)	(0.060)	(0.057)	(0.058)	(0.062)	(0.060)	(0.061)	(0.060)	(0.061)
Kaitz*2010	0.105	0.103*	0.098	0.117*	0.121*	0.116	0.112*	0.107*	0.133*	0.138*
	(0.071)	(0.062)	(0.063)	(0.070)	(0.072)	(0.076)	(0.063)	(0.064)	(0.076)	(0.077)
lambda		0.028	0.07*				0.031	0.067		
		(0.050)	(0.041)				(0.050)	(0.041)		
Observations	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148	1,148
R-squared	0.191			0.192	0.193	0.192			0.193	0.194
ll		2,057.721	2,059.021				2,058.200	2,059.391		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10) *** p<0.01, ** p<0.05, * p<0.1

Table A-13. Robustness check: Within Group Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 406 areas level (307 selected regions), 1997-2010, all regressions contain control variables, area and year effects..

	(1) FE	(2) SEMP Commuting	(3) SEMP Contiguity	(4) SLXP Commuting	(5) SLXP Contiguity	(6) FE	(7) SEMP Commuting	(8) SEMP Contiguity	(9) SLXP Commuting	(10) SLXP Contiguity
Kaitz Index	-0.032	-0,03	-0,032	-0.033	-0.033	-0.034	-0,032	-0,033	-0.035	-0.035
	(0.030)	(0.027)	(0.027)	(0.030)	(0.030)	(0.031)	(0.027)	(0.027)	(0.031)	(0.031)
w*Kaitz Index				0.054	0.023				0.063	0.028
				(0.117)	(0.090)				(0.118)	(0.090)
GVA						-0.134	-0,15	-0,136	-0.145	-0.138
						(0.187)	(0.159)	(0.156)	(0.189)	(0.188)
Share Public	0.004	0,004	0,003	0.004	0.004	0.004	0,004	0,004	0.004	0.004
	(0.028)	(0.025)	(0.025)	(0.028)	(0.028)	(0.028)	(0.025)	(0.025)	(0.028)	(0.028)
Kaitz*1999	-0.043*	-0,045	-0,044	-0.043*	-0.043*	-0.043*	-0,045	-0,044	-0.043*	-0.043*
	(0.023)	(0.029)	(0.028)	(0.023)	(0.023)	(0.023)	(0.029)	(0.028)	(0.023)	(0.023)
Kaitz*2000	0.002	0	0,002	0.002	0.002	0.003	0,001	0,002	0.003	0.003
	(0.020)	(0.028)	(0.028)	(0.020)	(0.020)	(0.021)	(0.028)	(0.028)	(0.021)	(0.020)
Kaitz*2001	0.005	0,003	0,004	0.005	0.005	0.009	0,007	0,008	0.009	0.009
	(0.021)	(0.027)	(0.027)	(0.021)	(0.021)	(0.022)	(0.028)	(0.027)	(0.022)	(0.022)
Kaitz*2002	0.033	0,031	0,032	0.033	0.033	0.038	0,037	0,037	0.038	0.038
	(0.022)	(0.027)	(0.027)	(0.022)	(0.022)	(0.023)	(0.028)	(0.028)	(0.023)	(0.023)
Kaitz*2003	0.049*	0,046*	0,049*	0.050*	0.049*	0.052*	0,049*	0,052*	0.053*	0.053*
	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)	(0.029)	(0.028)	(0.028)	(0.029)	(0.029)
Kaitz*2004	0.062**	0,059**	0,061**	0.062**	0.062**	0.066**	0,062**	0,065**	0.066**	0.066**
	(0.025)	(0.028)	(0.028)	(0.025)	(0.026)	(0.027)	(0.028)	(0.028)	(0.027)	(0.027)
Kaitz*2005	0.066**	0,063**	0,065**	0.065**	0.065**	0.070**	0,066**	0,068**	0.069**	0.069**
	(0.033)	(0.029)	(0.029)	(0.032)	(0.033)	(0.034)	(0.029)	(0.029)	(0.034)	(0.034)
Kaitz*2006	0.060*	0,057*	0,059**	0.059*	0.060*	0.063*	0,06**	0,062**	0.063*	0.063*
	(0.032)	(0.030)	(0.029)	(0.032)	(0.032)	(0.032)	(0.030)	(0.030)	(0.032)	(0.032)
Kaitz*2007	0.003	-0,002	0,002	0.002	0.003	0.006	0,001	0,004	0.005	0.005
	(0.028)	(0.029)	(0.029)	(0.028)	(0.028)	(0.029)	(0.029)	(0.029)	(0.029)	(0.029)
Kaitz*2008	-0.015	-0,02	-0,016	-0.016	-0.015	-0.011	-0,017	-0,013	-0.012	-0.012
	(0.034)	(0.029)	(0.028)	(0.034)	(0.034)	(0.035)	(0.029)	(0.028)	(0.035)	(0.035)
Kaitz*2009	-0.015	-0,016	-0,015	-0.016	-0.015	-0.012	-0,013	-0,012	-0.013	-0.012
	(0.039)	(0.029)	(0.029)	(0.039)	(0.039)	(0.040)	(0.029)	(0.029)	(0.040)	(0.040)
Kaitz*2010	0.083**	0,081***	0,082***	0.083**	0.083**	0.089**	0,087***	0,088***	0.088**	0.089**
	(0.040)	(0.030)	(0.029)	(0.040)	(0.040)	(0.042)	(0.030)	(0.030)	(0.042)	(0.042)
lambda		0,050*	0,016				0,055*	0,016		
		(0.029)	(0.023)				(0.029)	(0.023)		
Observations	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298	4,298
R-squared	0.091			0.091	0.091	0.091			0.091	0.091
ll		6,300.791	6,299.699				6,301.268	6,300.108		

(Robust) standard errors in parentheses (columns 1,4,5,6,9,10) *** p<0.01, ** p<0.05, * p<0.1

Table A-14. Robustness check: SGMM-SLXP Estimates of Minimum Wage Effects on Employment, 16 years to retirement age, 1997-2010, all regressions contain control variables, area and year effects.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions (74 selected regions)				138 regions (82 selected regions)				406 regions (307 selected regions)			
	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Contiguity	SGMM SLXP Contiguity	SGMM SLXP Commut.	SGMM SLXP Commut	SGMM SLXP Contiguity.	SGMM SLXP Contiguity	SGMM SLXP Commut.	SGMM SLXP Commut	SGMM SLXP Contiguity.	SGMM SLXP Contiguity
Kaitz Index	-0.008	-0.015	-0.091	-0.074	-0.025	-0.035	-0.069	-0.092	0.003	-0.025	-0.028	-0.049
	(0.070)	(0.074)	(0.062)	(0.076)	(0.071)	(0.071)	(0.070)	(0.076)	(0.045)	(0.045)	(0.044)	(0.048)
w*Kaitz Index	-0.368*	-0.386*	-0.015	-0.014	-0.276***	-0.280***	-0.119*	-0.116*	-0.203**	-0.231**	-0.018	-0.028
	(0.202)	(0.206)	(0.053)	(0.051)	(0.084)	(0.081)	(0.060)	(0.060)	(0.103)	(0.100)	(0.075)	(0.075)
E_{t-1}	0.204***	0.204***	0.194***	0.187***	0.142***	0.141***	0.119**	0.117*	0.219***	0.221***	0.216***	0.216***
	(0.058)	(0.058)	(0.059)	(0.058)	(0.050)	(0.050)	(0.058)	(0.060)	(0.029)	(0.029)	(0.029)	(0.029)
GVA		-0.156		0.304		-0.239		-0.458		-0.568**		-0.353
		(0.303)		(0.357)		(0.400)		(0.596)		(0.235)		(0.265)
Share Public	-0.254***	-0.247***	-0.306***	-0.299***	-0.200**	-0.198**	-0.368***	-0.379***	-0.201***	-0.194***	-0.213***	-0.208***
	(0.081)	(0.081)	(0.099)	(0.100)	(0.076)	(0.076)	(0.098)	(0.101)	(0.040)	(0.039)	(0.064)	(0.065)
Kaitz*1999	-0.045	-0.043	-0.035	-0.035	-0.015	-0.013	-0.032	-0.022	-0.061**	-0.048*	-0.061**	-0.052**
	(0.050)	(0.050)	(0.041)	(0.048)	(0.045)	(0.044)	(0.043)	(0.034)	(0.025)	(0.025)	(0.025)	(0.025)
Kaitz*2000	0.054	0.053	0.077**	0.072*	0.049	0.052	0.050	0.046	0.013	0.027	0.023	0.033
	(0.044)	(0.047)	(0.037)	(0.038)	(0.045)	(0.045)	(0.051)	(0.037)	(0.026)	(0.028)	(0.026)	(0.029)
Kaitz*2001	0.050	0.055	0.017	0.007	0.008	0.020	-0.020	-0.031	-0.014	0.013	-0.013	0.004
	(0.035)	(0.037)	(0.036)	(0.042)	(0.048)	(0.053)	(0.059)	(0.045)	(0.025)	(0.028)	(0.025)	(0.028)
Kaitz*2002	0.063	0.067	0.067**	0.052	0.008	0.021	-0.010	-0.000	0.008	0.040	0.017	0.038
	(0.044)	(0.046)	(0.032)	(0.038)	(0.047)	(0.055)	(0.053)	(0.046)	(0.028)	(0.032)	(0.028)	(0.032)
Kaitz*2003	0.112	0.121*	0.094	0.076	0.086*	0.098*	0.076	0.056	0.026	0.051	0.032	0.050
	(0.069)	(0.068)	(0.061)	(0.070)	(0.049)	(0.050)	(0.050)	(0.040)	(0.032)	(0.033)	(0.030)	(0.034)
Kaitz*2004	0.084	0.090	0.080	0.064	0.063	0.074	0.063	0.043	0.032	0.059*	0.053*	0.071**
	(0.058)	(0.058)	(0.049)	(0.055)	(0.059)	(0.063)	(0.069)	(0.051)	(0.028)	(0.030)	(0.028)	(0.031)
Kaitz*2005	0.093**	0.097**	0.078*	0.062	0.032	0.044	0.028	0.033	0.038	0.063	0.048	0.064
	(0.037)	(0.040)	(0.039)	(0.043)	(0.054)	(0.059)	(0.067)	(0.048)	(0.036)	(0.038)	(0.036)	(0.039)
Kaitz*2006	0.108*	0.115*	0.115**	0.099	-0.003	0.010	-0.027	-0.027	0.038	0.060	0.026	0.039
	(0.062)	(0.062)	(0.051)	(0.061)	(0.074)	(0.076)	(0.069)	(0.050)	(0.042)	(0.044)	(0.043)	(0.045)
Kaitz*2007	-0.002	-0.002	0.004	-0.009	0.003	0.015	-0.013	-0.006	-0.025	-0.003	-0.021	-0.006
	(0.066)	(0.069)	(0.050)	(0.050)	(0.054)	(0.057)	(0.052)	(0.037)	(0.035)	(0.037)	(0.035)	(0.038)
Kaitz*2008	0.056	0.066	0.063	0.041	0.069	0.084	0.018	-0.002	-0.034	-0.009	-0.010	0.008
	(0.058)	(0.061)	(0.055)	(0.067)	(0.053)	(0.057)	(0.058)	(0.051)	(0.038)	(0.041)	(0.037)	(0.041)
Kaitz*2009	0.021	0.027	0.052	0.033	-0.058	-0.042	-0.032	-0.064	-0.028	-0.004	-0.021	-0.006
	(0.061)	(0.065)	(0.050)	(0.063)	(0.077)	(0.080)	(0.074)	(0.055)	(0.046)	(0.048)	(0.045)	(0.047)
Kaitz*2010	0.082	0.087	0.123**	0.104*	0.020	0.041	0.049	-0.007	0.076*	0.112**	0.101**	0.126***
	(0.075)	(0.080)	(0.059)	(0.061)	(0.080)	(0.091)	(0.095)	(0.058)	(0.043)	(0.046)	(0.044)	(0.048)
Observations	962	962	962	962	1,066	1,066	1,066	1,066	3,991	3,991	3,991	3,991
No. of instruments	70	71	70	71	70	71	70	71	70	71	70	71

*** p<0.01, ** p<0.05, * p<0.1. Test statistics are provided in the next table.

Table A-15. Robustness check: SGMM-SLXP Estimates of Minimum Wage Effects on Employment, test statistics of estimates in table A-14.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions (74 selected regions)				138 regions (82 selected regions)				406 regions (307 selected regions)			
	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA
Number of instruments	70	71	70	71	70	71	70	71	70	71	70	71
Arellano-Bond test for AR in first differences												
AR(1)	-4.6839	-4.7036	-4.7697	-4.7434	-5.1199	-5.1122	-4.8914	-4.8467	-12.7983	-12.7433	-12.7322	-12.6661
Prob > z	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
AR(2)	0.1172	0.1139	0.0270	0.0273	-2.1746	-2.1938	-2.3382	-2.3807	-1.1360	-1.1257	-1.1810	-1.1811
Prob > z	0.9067	0.9093	0.9785	0.9783	0.0297	0.0283	0.0194	0.0173	0.2560	0.2603	0.2376	0.2376
Hansen test of overidentified restrictions												
J	37.7582	38.0452	33.9748	33.3067	43.2643	43.0232	43.3846	43.0660	31.0707	31.9712	32.5959	34.1737
Prob> chi2	0.3889	0.3764	0.5652	0.5974	0.1889	0.1958	0.1855	0.1946	0.7020	0.6607	0.6313	0.5557
Difference-in-Sargan tests of exogeneity of instrument subsets												
- GMM instruments -												
C	36.3796	36.7445	31.1543	30.6598	42.8531	42.7874	42.3823	42.2339	30.7967	31.7519	31.9437	33.8926
Prob > chi2	0.3142	0.2995	0.5592	0.5841	0.1170	0.1184	0.1269	0.1302	0.5772	0.5292	0.5196	0.4243
- Instrumented variables in levels and first differences -												
C	21.5235	21.5892	20.6872	19.2652	24.1955	23.5716	23.2064	21.5415	15.3938	15.1661	13.0643	12.9796
Prob > chi2	0.2538	0.2010	0.2955	0.3135	0.1487	0.1315	0.1828	0.2030	0.6348	0.5835	0.7877	0.7376
- Instrumented variables only in levels -												
C	24.2285	24.7322	22.4313	21.8255	36.0395	35.2394	34.3011	34.1209	17.1602	18.1336	19.0155	19.2990
Prob > chi2	0.4486	0.4204	0.5535	0.5897	0.0544	0.0649	0.0795	0.0825	0.8417	0.7965	0.7512	0.7359

Table A-16. Robustness check: SGMM-SLXP Estimates of Minimum Wage Effects on Employment, excluding direct effect of Minimum Wage, 16 years to retirement age, 1997-2010, all regressions contain control variables, area and year effects..

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions (74 selected regions)				138 regions (82 selected regions)				406 regions (307 selected regions)			
	SGMM SLXP Commut.	SGMM SLXP Commut.	SGMM SLXP Contiguit y	SGMM SLXP Contiguit y	SGMM SLXP Commut.	SGMM SLXP Commut	SGMM SLXP Contiguit y.	SGMM SLXP Contiguit y	SGMM SLXP Commut.	SGMM SLXP Commut	SGMM SLXP Contiguit y.	SGMM SLXP Contiguit y
Kaitz Index	-	-	-	-	-	-	-	-	-	-	-	-
	-	-	-	-	-	-	-	-	-	-	-	-
w*Kaitz Index	-0.390**	-0.410**	-0.031	-0.025	-0.279***	-0.288***	-0.184***	-0.128***	-0.202**	-0.127	0.026	0.016
	(0.185)	(0.188)	(0.026)	(0.025)	(0.069)	(0.069)	(0.062)	(0.038)	(0.099)	(0.092)	(0.051)	(0.056)
E_{t-1}	0.207***	0.205***	0.205***	0.204***	0.128**	0.128**	0.129***	0.119**	0.219***	0.204***	0.204***	0.203***
	(0.059)	(0.060)	(0.045)	(0.045)	(0.049)	(0.049)	(0.048)	(0.048)	(0.028)	(0.028)	(0.027)	(0.027)
GVA		-0.130		0.145		-0.256		-0.386		-0.750***		-0.521**
		(0.288)		(0.311)		(0.368)		(0.477)		(0.220)		(0.206)
Share Public	-0.244***	-0.239***	-0.243***	-0.244***	-0.195***	-0.195***	-0.222***	-0.248**	-0.201***	-0.198***	-0.217***	-0.209***
	(0.084)	(0.085)	(0.074)	(0.076)	(0.073)	(0.073)	(0.080)	(0.124)	(0.039)	(0.062)	(0.060)	(0.034)
Kaitz*1999	-0.047	-0.049	-0.075*	-0.071*	-0.019	-0.022	-0.055	-0.022	-0.059*	-0.028	-0.035	-0.030
	(0.047)	(0.049)	(0.040)	(0.040)	(0.052)	(0.051)	(0.054)	(0.064)	(0.031)	(0.032)	(0.031)	(0.027)
Kaitz*2000	0.054	0.046	0.029	0.031	0.056	0.053	0.002	0.038	0.015	0.041	0.037	0.043
	(0.052)	(0.057)	(0.041)	(0.041)	(0.059)	(0.056)	(0.058)	(0.049)	(0.033)	(0.037)	(0.035)	(0.031)
Kaitz*2001	0.053	0.051	-0.029	-0.027	0.017	0.022	-0.039	-0.012	-0.011	0.040	0.023	0.034
	(0.059)	(0.062)	(0.045)	(0.045)	(0.062)	(0.062)	(0.063)	(0.067)	(0.034)	(0.039)	(0.036)	(0.032)
Kaitz*2002	0.066	0.061	0.008	0.004	0.022	0.029	-0.048	0.009	0.010	0.068	0.043	0.052
	(0.067)	(0.071)	(0.052)	(0.051)	(0.066)	(0.068)	(0.068)	(0.065)	(0.041)	(0.046)	(0.042)	(0.037)
Kaitz*2003	0.113*	0.114*	0.061	0.059	0.088	0.091*	0.028	0.073	0.029	0.075**	0.061*	0.071**
	(0.064)	(0.065)	(0.044)	(0.045)	(0.054)	(0.052)	(0.054)	(0.055)	(0.030)	(0.035)	(0.035)	(0.028)
Kaitz*2004	0.083	0.081	0.030	0.027	0.076	0.078	0.014	0.084	0.034	0.090**	0.070*	0.083***
	(0.062)	(0.064)	(0.044)	(0.045)	(0.066)	(0.066)	(0.066)	(0.058)	(0.036)	(0.039)	(0.036)	(0.031)
Kaitz*2005	0.097	0.093	0.042	0.041	0.049	0.054	-0.018	0.030	0.041	0.102**	0.082**	0.086**
	(0.068)	(0.072)	(0.053)	(0.052)	(0.068)	(0.067)	(0.065)	(0.063)	(0.044)	(0.045)	(0.042)	(0.037)
Kaitz*2006	0.107	0.107	0.038	0.040	0.008	0.012	-0.065	-0.003	0.041	0.101*	0.085	0.081
	(0.065)	(0.068)	(0.046)	(0.047)	(0.080)	(0.079)	(0.079)	(0.084)	(0.054)	(0.058)	(0.056)	(0.053)
Kaitz*2007	-0.004	-0.011	-0.065	-0.065	0.011	0.015	-0.058	-0.004	-0.022	0.030	0.017	0.012
	(0.067)	(0.071)	(0.064)	(0.063)	(0.073)	(0.071)	(0.071)	(0.064)	(0.044)	(0.044)	(0.042)	(0.037)
Kaitz*2008	0.057	0.058	0.018	0.018	0.072	0.080	0.003	0.013	-0.032	0.019	0.002	0.012
	(0.062)	(0.063)	(0.048)	(0.049)	(0.070)	(0.068)	(0.068)	(0.067)	(0.043)	(0.041)	(0.040)	(0.036)
Kaitz*2009	0.021	0.018	-0.013	-0.013	-0.047	-0.038	-0.123	-0.074	-0.025	0.046	0.029	0.045
	(0.074)	(0.078)	(0.048)	(0.050)	(0.097)	(0.094)	(0.097)	(0.086)	(0.053)	(0.059)	(0.057)	(0.051)
Kaitz*2010	0.082	0.077	0.040	0.037	0.032	0.048	-0.050	0.029	0.079	0.157***	0.124**	0.145***
	(0.081)	(0.083)	(0.062)	(0.064)	(0.100)	(0.102)	(0.103)	(0.099)	(0.049)	(0.053)	(0.050)	(0.044)
Observations	962	962	962	962	1,066	1,066	1,066	1,066	3,991	3,991	3,991	3,991
Number of instruments	69	70	69	70	69	70	81	70	69	70	69	70

*** p<0.01, ** p<0.05, * p<0.1. Test statistics are provided in the next table.

Table A-17. Robustness check: SGMM-SLXP Estimates of Minimum Wage Effects on Employment, excluding direct effect of Minimum Wage, test statistics of estimates in table A-16.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	140 regions				138 regions				406 regions			
	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA	SGMM SLXP Commut.	SGMM SLXP Commut. GVA	SGMM SLXP Contiguity	SGMM SLXP Contiguity GVA
Number of instruments	69	70	69	70	69	70	81	70	69	70	69	70
Arellano-Bond test for AR in first differences												
AR(1)	-4.7113	-4.7271	-4.7734	-4.7995	-5.1369	-5.1179	-5.1882	-5.0785	-12.7176	-12.7520	-12.7855	-12.9756
Prob > z	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
AR(2)	0.1319	0.1211	0.1731	0.1963	-2.2395	-2.2552	-2.1693	-2.2706	-1.1354	-1.5344	-1.5582	-1.6148
Prob > z	0.8951	0.9036	0.8626	0.8444	0.0251	0.0241	0.0301	0.0232	0.2562	0.1249	0.1192	0.1063
Hansen test of overidentified restrictions												
J	37.4905	37.9077	34.7615	34.5127	44.2721	44.2255	44.5016	43.0701	31.0790	38.6340	36.9930	38.3717
Prob> chi2	0.4006	0.3824	0.5274	0.5394	0.1620	0.1632	0.1563	0.1944	0.7016	0.3515	0.4229	0.3625
Difference-in-Hansen tests of exogeneity of instrument subsets												
- GMM instruments for levels -												
C	36.4761	37.0538	29.2444	29.1380	43.4445	43.6545	43.6167	41.8412	30.8432	38.4178	36.7723	34.9679
Prob > chi2	0.3102	0.2873	0.6547	0.6599	0.1055	0.1016	0.1023	0.1391	0.5749	0.2374	0.2983	0.3747
- Instrumented variables in levels and first differences -												
C	21.5513	21.6251	19.7974	17.6701	25.7586	25.9758	26.0593	27.5163	15.6797	20.6276	20.6685	17.3545
Prob > chi2	0.3072	0.2491	0.4069	0.4776	0.1371	0.1003	0.1285	0.0698	0.6785	0.2986	0.3554	0.4989
- Instrumented variables only in levels -												
C	23.2392	23.8295	19.7863	19.6524	36.7602	36.5592	36.8522	31.3911	17.1580	28.8411	25.7375	27.6381
Prob > chi2	0.5057	0.4714	0.7089	0.7164	0.0462	0.0484	0.0453	0.1428	0.8418	0.2262	0.3666	0.2756

Table A-18. Moran's *I* statistics for the residuals of the FE and commuting weights matrix based regression models in Tables 1 to 6

Columns in Tables 1-6		(1)	(2)	(4)	(6)	(7)	(9)
Geography / Model		FE	SEMP Commuting Matrix	SLXP Commuting Matrix	FE GVA	SEMP Commuting Matrix GVA	SLXP Commuting Matrix GVA
140 regions, without yearly iDiD terms	*I*	0.027	0.028	0.027	0.004	0.005	0.005
	I/V	2.484	2.567	2.504	0.405	0.415	0.432
140 regions, with yearly IDiD terms	*I*	0.009	0.010	0.010	-0.003	-0.004	-0.003
	I/V	0.847	0.938	0.881	-0.305	-0.323	-0.269
138 regions, without yearlyIDiD terms	*I*	0.036	0.037	0.036	0.024	0.025	0.024
	I/V	3.313	3.424	3.271	2.239	2.323	2.204
138 regions, with yearlyIDiD terms	*I*	0.029	0.031	0.029	0.020	0.021	0.020
	I/V	2.696	2.871	2.685	1.824	1.930	1.806
406 regions, without yearlyIDiD terms	*I*	0.401	0.401	0.401	0.396	0.396	0.396
	I/V	50.841	50.861	50.830	50.127	50.159	50.119
406 regions, without yearly IDiD terms	*I*	0.396	0.396	0.396	0.393	0.393	0.393
	I/V	50.120	50.206	50.121	49.769	49.836	49.769

Table A-19. Moran's *I* statistics for the residuals of the FE and contiguity weights matrix based regression models in Tables 1 to 6

Columns in Tables 1-6		(1)	(2)	(4)	(6)	(7)	(9)
Geography / Model		FE	SEMP Contiguity Matrix	SLXP Contiguity Matrix	FE GVA	SEMP Contiguity Matrix GVA	SLXP Contiguity Matrix GVA
140 regions, without yearly IDiD terms	*I*	0.021	0.021	0.020	0.006	0.006	0.006
	I/V	1.540	1.578	1.533	0.473	0.476	0.474
140 regions, with yearly IDiD terms	*I*	0.001	0.001	0.000	-0.003	-0.004	-0.004
	I/V	0.057	0.071	-0.014	-0.261	-0.272	-0.319
138 regions, without yearlyIDiD terms	*I*	0.037	0.037	0.038	0.029	0.029	0.031
	I/V	3.042	3.093	3.143	2.386	2.405	2.522
138 regions, with yearlyIDiD terms	*I*	0.030	0.031	0.031	0.024	0.025	0.025
	I/V	2.441	2.556	2.523	1.971	2.042	2.080
406 regions, without yearlyIDiD terms	*I*	0.030	0.030	0.030	0.023	0.023	0.023
	I/V	4.546	4.572	4.579	3.450	3.464	3.518
406 regions, without yearly IDiD terms	*I*	0.020	0.020	0.020	0.017	0.017	0.017
	I/V	3.010	3.055	3.059	2.574	2.598	2.636

Table A-20. Moran's *I* statistics for the residuals of the SGMM SLXP regression models in Tables 7, A-4, and A-6.

		140 regions		138 regions		406 regions	
		SGMM SLXP	SGMM SLXP GVA	SGMM SLXP	SGMM SLXP GVA	SGMM SLXP	SGMM SLXP GVA
Columns in table 7		(1)	(2)	(3)	(4)	(5)	(6)
Models without a direct Kaitz lag term, based on commuting matrix	*I*	0.070	0.038	0.099	0.068	0.459	0.458
	I/V	6.215	3.346	8.740	5.990	56.033	55.928
Columns in table A-6		(1)	(2)	(3)	(4)	(5)	(6)
Models without a direct Kaitz term, based on contiguity matrix	*I*	0.081	0.087	0.108	0.077	0.199	0.203
	I/V	5.807	6.307	8.598	6.111	29.421	29.899
Columns in table A-4		(1)	(2)	(5)	(6)	(9)	(10)
Models with a direct Kaitz lag term, based on commuting matrix	*I*	0.071	0.068	0.099	0.095	0.459	0.458
	I/V	6.232	5.974	8.771	8.404	56.043	55.937
Columns in table A-4		(3)	(4)	(7)	(8)	(11)	(12)
Models with a direct Kaitz lag term, based on contiguity matrix	*I*	0.132	0.132	0.154	0.088	0.198	0.199
	I/V	9.490	9.522	12.246	7.024	29.267	29.293

Table A-21. Pesaran's *CD* statistics for the residuals of the FE and commuting weights matrix based regression models in Tables 1 to 6

Columns in Tables 1-6	(1)	(3)	(5)	(6)	(8)	(10)
	FE	SEMP Commutin g Matrix	SLXP Commutin g Matrix	FE GVA	SEMP Commutin g Matrix GVA	SLXP Commutin g Matrix GVA
140 regions, without yearly IDiD terms	0.497	0.709	0.410	-1.326	-1.285	-1.431
140 regions, with yearly IDiD terms	0.223	0.232	0.044	-1.516	-1.526	-1.661
138 regions, without yearly IDiD terms	-0.186	0.502	0.426	-0.997	-0.460	-0.569
138 regions, with yearly IDiD terms	-0.518	-0.179	0.100	-1.421	-1.156	-0.977
406 regions, without yearly IDiD terms	-1.476	-1.478	-1.248	-2.259	-2.202	-2.145
406 regions, with yearly IDiD terms	-1.243	-1.220	-1.140	-2.038	-1.967	-1.962

Table A-22. Pesaran's *CD* statistics for the residuals of the FE and Contiguity weights matrix based regression models in Tables 1 to 6

Columns in Tables 1-6	(1)	(3)	(5)	(6)	(8)	(10)
	FE	SEMP Contiguity Matrix	SLXP Contiguity Matrix	FE GVA	SEMP Contiguity Matrix GVA	SLXP Contiguity Matrix GVA
140 regions, without yearly IDiD terms	0.497	0.703	0.682	-1.326	-1.278	-1.256
140 regions, with yearly IDiD terms	0.223	0.265	0.500	-1.516	-1.522	-1.336
138 regions, without yearly IDiD terms	-0.186	0.067	-0.165	-0.997	-0.825	-0.851
138 regions, with yearly IDiD terms	-0.518	-0.341	-0.515	-1.421	-1.304	-1.249
406 regions, without yearly IDiD terms	-1.476	-1.467	-1.235	-2.259	-2.248	-2.093
406 regions, with yearly IDiD terms	-1.243	-1.247	-1.045	-2.038	-2.039	-1.826

Table A-23. Pesaran's *CD* statistics for the residuals of the SGMM SLXP regression models in Tables 7, A-4, and A-6.

	140 regions		138 regions		406 regions	
	SGMM SLXP	SGMM SLXP GVA	SGMM SLXP	SGMM SLXP GVA	SGMM SLXP	SGMM SLXP GVA
Columns in table 7	(1)	(2)	(3)	(4)	(5)	(6)
Models without a direct Kaitz lag term, based on commuting matrix	-1.812	-2.153	-1.898	-1.339	0.502	0.577
Columns in table A-6	(1)	(2)	(3)	(4)	(5)	(6)
Models without a direct Kaitz term, based on contiguity matrix	-1.733	-1.572	-1.282	-1.781	0.618	0.024
Columns in table A-4	(1)	(2)	(5)	(6)	(9)	(10)
Models with a direct Kaitz lag term, based on commuting matrix	-1.811	-1.865	-1.955	-2.010	0.491	0.589
Columns in table A-4	(3)	(4)	(7)	(8)	(11)	(12)
Models with a direct Kaitz lag term, based on contiguity matrix	-1.940	-2.023	-1.779	-1.594	0.572	0.774